Essays on sovereign bond pricing and inflation-linked products

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Contents

Introduction

This doctoral dissertation consists of three chapters on the pricing of sovereign debt and inflation-linked products. The first chapter examines the relative pricing of nominal and inflation-linked debt of the three largest Eurozone sovereign issuers. Its main contribution is to present evidence of a selective default premium in real bond yields. The second chapter shifts its focus to the US inflation-linked product markets and quantifies liquidity premium in TIPS and inflation swap rates. The size of this compensation for exposure to asset level and liquidity risk helps to explain a large part of the TIPS-Treasury puzzle. The third chapter studies whether nominal bond markets are segmented across different maturities and contributes to the policy discussion on long term discount rates of the Solvency II Directive.

Sovereign bonds and inflation-linked products are crucially important financial instruments for a wide range of large institutional investors, with a special emphasis on pension and insurance funds. The inclusion of inflation-linked assets in investment portfolios facilitates hedging against inflation risk and the indexation of long term liabilities. Sovereign bonds also pay a major role in both sides of the balance sheet: long maturity nominal bonds often serve as an input to attain precise estimates of long term discount rates for asset management and for valuation of liabilities for regulatory purposes. Additionally, the adequate understanding of the risk profile of sovereign debt is crucial not only from a risk management perspective, but also from a monetary policy point of view. By identifying the risk premiums in the yields of these securities, institutions can better manage their portfolios and comply with prudential regulation, whereas governments can issue bonds that are correctly priced.

The first chapter[1] presents evidence of a selective default risk premium in inflation-linked sovereign bond (ILB) yields of Germany, France and Italy. Selective default is an event, in which a sovereign issuer chooses not to meet obligations on a class of bonds, while servicing her other debt. This effect is identified by means of a unique empirical strategy.

[1]This chapter is based on the working paper titled: “Not risk free: The relative pricing of euro area inflation-indexed and nominal bonds”.

First, we construct breakeven yields, the difference between ILB and nominal yields, from maturity-matched bond pairs within each country. In the next step, we pair these breakeven rates across countries, by minimizing maturity gaps between the original bond pairs. The result is the spread on breakeven strategy, which is the difference between two bond pairs from two different countries. The differencing controls for common Eurozone level components in yields, such as the effect of inflation expectations, monetary policy or interest rate risk. What the differencing does not take out is the exposure to risks that do not affect nominal and inflation-linked bonds equally within a country. We show that there are two systematic risk factors that drive a wedge between inflation expectations and the breakeven rate: liquidity and sovereign credit risks. This implies that yields of ILBs and nominal bonds carry different levels of liquidity and sovereign risk premia. The latter suggests that even without explicit seniority between the two types of bonds, the market fears that an issuer is more likely to selectively default on its riskier, inflation-linked debt in periods of financial distress. Our findings are also linked to the ILB-nominal puzzle of Fleckenstein et al. (2014). In a frictionless world, one can replicate a nominal bond with a portfolio of an ILB and inflation swap contracts. They find that the replicating portfolio has a lower price than the nominal bond, suggesting that ILBs are underpriced. We provide evidence that this underpricing is in part due to relative risk premium differences between nominal and inflation-linked debt: ILBs are less liquid, moreover investors perceive them to have higher credit risk during the financial and euro crises, further increasing the yield difference between the two securities.

The second chapter[2] examines the US Treasury bond and inflation swaps markets. We provide evidence that in both index-linked bond markets and inflation swap markets liquidity is an important determinant of prices. To study this phenomenon, we propose an asset pricing model with a liquidity risk factor and asset-specific liquidity characteristics. To estimate the effect of liquidity risk, we measure an assets exposure to a non-traded liquidity factor, which is derived from the measures of Amihud (2002) and Roll (1984). In addition, the level of liquidity is proxied by asset-level characteristics, following Krishnamurthy (2002) and Houweling et al. (2005). We conduct our analyses based on US data and under the assumption of either end of the spectrum: completely segmented or integrated markets. In our benchmark specifications, assuming segmentation, we find strong evidence that the level of liquidity, in contrast to liquidity risk, affects yields on inflation-indexed bonds, whereas inflation swap yields include a liquidity risk premium. We also quantify liquidity effects in nominal bond yields and find a small liquidity risk premium. These results are robust to the inclusion of various controls and to shifting to the proposition of integrated markets. Our second main contribution is that we examine

[2]This chapter is based on a joint working paper with Joost Driessen and Theo Nijman, titled: "The missing piece of the puzzle: Liquidity premiums in inflationindexed markets".

whether the above diversity in exposures to liquidity and liquidity risk could explain the persistent difference in relative bond prices, as documented in Fleckenstein et al. (2014). They show that there exists a substantial price difference between a nominal Treasury bond and its replicating portfolio that consists of a TIPS issue and inflation swap contracts. We provide evidence that a large part of the TIPS underpricing disappears when we control for the estimated liquidity effects in TIPS yields and inflation swaps rates.

The third chapter[3] provides comprehensive evidence on the pricing differences of short and long maturity nominal bonds. Long maturity bonds are popular assets among investors with long investment horizon, such as pension funds and insurance companies. Despite its practical importance and potential welfare consequences, modelling and examining the long end of the nominal term structure has attracted little attention in the academic literature. This chapter aims to fill this gap by studying the differential pricing of short and long maturity bonds, especially focusing on segmentation in yields and liquidity. By using data on German nominal bonds between 2005 and 2015, we aim is to answer the following question: Are yields of long-maturity bonds distorted by demand pressure of clientele investors, regulatory effects, or default, flight-to-safety or liquidity premiums? We find that although there are statistically significant differences in the pricing and drivers of short and long maturity bonds, the corresponding economic effects are rather small. This means that long yields are not extensively distorted by demand pressure, default or liquidity premiums, therefore there is little evidence for substantial yield segmentation. Additionally, we present evidence for some degree of liquidity segmentation across short and long maturities, with equally small economic effects. These two findings have important policy implications for the European insurance and pension regulatory framework, the Solvency II Directive. Part of this discussion on valuation of pension and insurance liabilities is on the modelling of long term discount rates. The current approach is based on the ultimate forward rate method, an extrapolation technique used to calculate discount rates for maturities beyond the last liquid point, based on statistical models and interest rate swaps. However, in light of our empirical result based on a simple method for extrapolation, this practice seems unnecessary: if long maturity bond yields are not distorted, we could extrapolate long term discount rates from these yields observed in bonds markets.

[3]This chapter is based on a joint working paper with Joost Driessen and Theo Nijman, titled: “Much ado about nothing: A study of differential pricing and liquidity of short and long term bonds”.

Chapter 1

Not risk free: The relative pricing of euro area inflation-indexed and nominal bonds

1.1 Introduction

Understanding the relative pricing of inflation-linked and nominal sovereign debt is important. First, these securities directly determine the breakeven inflation rate, the yield difference between nominal and inflation-linked bonds, henceforth ILBs, which is a market-based proxy for inflation expectations. However, if different levels of risk premia drove these bond prices, the breakeven rate would be distorted. Consistent with this idea, Pflueger and Viceira (2015) and Driessen et al. (2014) show that the liquidity premium differs among indexed and nominal bonds. Second, these securities play an important role in the portfolio choice of a wide range of investors. For instance, pension funds and insurers are seeking inflation-linked products, thus indexed-bonds too, to incorporate into their portfolios. Moreover, the adequate understanding of the risk profile of sovereign bonds is crucial not only from the risk management perspective of investors, but also from a monetary policy point of view. By identifying the risk premia in the

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yields of these securities, institutions can better manage their portfolios and comply with prudential regulation, whereas governments can issue bonds that are correctly priced.

The key result of this paper is the empirical evidence of selective default risk premium in inflation-linked sovereign bond yields of Germany, France and Italy. We define selective default as an event in which a sovereign issuer chooses not to meet obligations on a class of bonds, while servicing its other debt. We identify this effect from the difference of breakeven rates from pairs of countries. Differencing eliminates common components, such as the effect of inflation expectations, monetary policy or interest rate risk. What the differencing does not take out is the exposure to risks that do not affect nominal and inflation-linked bonds equally within a country. We show that there are two systematic risk factors that drive a wedge between inflation expectations and the breakeven rate: these are liquidity and sovereign credit risks. This implies that yields of ILBs and nominal bonds carry different levels of liquidity and sovereign risk premia. The latter suggests that even without explicit seniority between the two types of bonds, the market perceives that an issuer is more likely to selectively default on its riskier, inflation-linked debt in periods of financial distress.

The idea of comparing yields of securities with similar exposures to certain risks is not new in the literature. Longstaff (2004) compares yields of US Treasuries to those of bonds issued by the Refcorp (Resolution Funding Corporation), whereas Schwarz (2015) examines yield differences of German federal government bonds and bonds issued by KFW, a government owned development bank. The key feature of theses agency bonds is that they have explicit government guarantees, and consequently the same credit risk as government bonds. However, the liquidity of government bonds is substantially higher and thus the yield difference measures general market liquidity conditions. What we do in this paper is similar but goes the other way around: while controlling for liquidity on both the nominal and inflation-linked bond markets the same way, we show that the remaining yield difference is attributed to sovereign risk. This idea is also consistent with the alternative interpretation of the Refcorp and KfW spreads - some say that these yield differentials, rather than capturing liquidity, can also be interpreted as breakup or selective default risk measures.

The secondary contribution of this paper is to provide partial explanation for the ILB-nominal puzzle documented by Fleckenstein et al. (2014). They claim that there exists a persistent mispricing between nominal bonds and ILBs of the US and other G7 countries on a significant scale. In a frictionless world, one can replicate a nominal bond with a portfolio of an ILB and inflation swaps. They find that the replicating portfolio has a lower price than the nominal bonds, suggesting that ILBs are underpriced. For the US market, Driessen et al. (2014) show that a large part of this price discrepancy is attributable to liquidity risk. However, there are other factors that could drive the mispricing, namely the impact of the deflation option[1] embedded in ILBs, liquidity and counterparty risk premia in the inflation swap quotes, or even different levels of selective default risk premia in nominal and real bonds.

The fore mentioned identification strategy can also be derived from the ILB-nominal puzzle, substituting the breakeven rates by the mispricing between nominal bonds and their replicating portfolios. Instead of examining these two prices in one country, we take this price difference and compare across countries. A unique feature of this cross-country sample is that in these three euro area countries both inflation swaps and inflation-indexed bonds are linked to the same price index[2] and the same deflation protection applies to all bonds. Therefore, as a result of the differencing, the swap component and the price effect of the deflation option mutually cancel out, reducing the new strategy to four bonds or a spread on two breakeven rates. The differencing sheds light on the drivers of the ILB-nominal puzzle: inflation swap quotes or the value of the deflation option cannot account for the overall magnitude of the puzzle. Second, we estimate the difference in liquidity and credit premia in ILB and nominal bonds and find that although the mean effect of liquidity is small, these two effects can explain the persistent nature of the puzzle. And lastly, we find that investors perceived ILBs to have higher sovereign risk exposure than nominal bonds during the financial and euro crises, further increasing the yield difference between the two securities.

Unlike most papers in the literature, we do not restrict our attention to examining the nominal sovereign spread. Our primary aim is to understand what drives the wedge between breakeven rates and inflation expectations, in other words the relative pricing of indexed and nominal sovereign bonds. Other papers looking at the relative pricing of nominal and indexed sovereign bonds are Campbell et al. (2009), Christensen and Gillan (2011), Pflueger and Viceira (2011, 2015), Fleckenstein et al. (2014), Fleckenstein (2013) and Driessen et al. (2014). Fleckenstein (2013) specifically focuses on the relative pricing of nominal and indexed bonds in G7 countries, whereas Driessen et al. (2014) show that most of the price difference between nominal and indexed US Treasuries is due to liquidity risk premium in prices.

By exploring the liquidity features of indexed and nominal sovereign bonds, we contribute to the long-standing literature on the effect of liquidity on asset prices (Amihud and Mendelson (1986); Amihud (2002); Bekaert et al. (2007) among many others). More

[1]The deflation floor provides protection for investors when negative inflation occurs. In the absence of the par floor, negative inflation would erode the value of the principal. In all European inflation-linked bonds, the principal value is protected against deflation, but not the intermediate coupon payments.

[2]HICP stands for Harmonized Index of Consumer Prices. This index is the weighted average of inflation of Eurozone countries and is published by the European Central Bank on a monthly basis.

specifically, we provide new evidence of liquidity risk being priced in major Eurozone sovereign bond markets. Other studies often examine liquidity in the context of spillover effects between European nominal sovereign bond and CDS markets (Calice et al., 2011) or focus on specific markets to show how liquidity improved upon EBC interventions (Pelizzon et al., mingb). Moreover, Darbha and Dufour (2014) show that even after controlling for interest rate and credit risks similarly to Fama and French (1993), liquidity is an important determinant of sovereign yields both in the cross-section and during the financial crisis.

Naturally, this paper also links to the strand of literature on European nominal sovereign market. Ejsing et al. (2012) investigate the dynamics of credit risk premium in bank and sovereign CDSs during the financial crisis, especially focusing on the effect of government rescue packages. Moreover, papers also examine the information content of sovereign CDS contracts and bonds (for instance Fontana and Scheicher (2010)) or look at the basis, the yield difference between these two assets (Arce et al., 2011; Palladini and Portes, 2011). In this paper there is novel evidence on the price of credit risk on both nominal and inflation linked sovereign bonds. Further, we also provide evidence on a subtler aspect of credit risk, namely selective default risk in the bonds under examination.

Our analysis is closest related to recent work on Euro area government bonds research that considers both liquidity and credit risks. Beber et al. (2009) disentangle the effects of liquidity and credit quality in 10 Eurozone countries to identify flight to quality and liquidity episodes. They show that liquidity is a non-trivial determinant of yields with an increasing prominence when flights occur, whereas credit quality affects valuation. On the other hand, by means of market related measures, Schwarz (2015) separates the components of yields due to liquidity and credit risk. She estimates a model of liquidity risk and finds that liquidity is priced in the cross-section of (nominal) sovereign debt. Ejsing et al. (2012) quantify liquidity and credit risk premia in German and French government bond yields based on a state-space model with two latent factors. Bai et al. (2012) examine what caused the sovereign debt crisis – illiquidity of markets or deteriorating credit conditions – and find spillover, but not feedback effect between aggregate level credit and liquidity risk in a country. And finally, Darbha and Dufour (2014) study the term structure of default and illiquidity in a sample of nominal Euro area government bonds, whereas Monfort and Renne (2014) present an arbitrage-free model of euro-area bond spreads, whose dynamics are driven by liquidity and credit risk. They find a non-diversifiable euro-area credit component in these yields.

Our work differs from the above papers in two main aspects. First, we examine a cross-country sample that goes beyond the nominal segment of bond markets and allows me to look at the relative pricing of nominal and inflation-indexed bonds in the Eurozone

– allowing me to set up clean, more stringent tests: the asset pricing tests we run are particularly strong, in the sense that we control for many confounding factors by the differencing. Therefore, in the subsequent step we are less likely to capture the effect of factors other than differential liquidity or credit risk. Second, the unique identification strategy based on differencing also allows me to address the empirical challenge of disentangling alternative explanations of the ILB-nominal puzzle.

The remainder of the chapter is organized as follows. Section 1.2 discusses the European bond markets and the methodology, whereas Section 1.3 explains the data and the estimation strategy. In Section 1.4 we present the empirical findings alongside with a discussion, and finally; Section 1.5 discusses possible extensions and concludes.

1.2 Are liquidity and credit risks priced in European nominal and inflation-indexed bonds?

In this section we shortly present the three major European sovereign bond markets: France, Germany and Italy. We specifically focus on market conventions, microstructure similarities and the inflation-linked bond segment. After showing why this is an ideal setting to study the relative pricing of real and nominal bonds, we discuss the identification strategy that helps to disentangle price effects of liquidity and credit risks in the corresponding bond yields. We present a multifactor asset pricing model with illiquidity and credit risk factors, inspired by Pastor and Stambaugh (2003), Acharya and Pedersen (2005), and Fama and French (1993). Then we show how to generalize this relationship to breakeven rates, and Appendix 1.A shows how this relates to the trading rule of Fleckenstein et al. (2014). The economic interpretation of the generalized model is the cornerstone of this paper: both liquidity and (sovereign) credit risk premia can differ among nominal and inflation-linked bonds of the same issuer.[3]

1.2.1 European bond markets

France, Germany and Italy are the three largest sovereign debt issuers of the Eurozone. These countries are part of a monetary union, consequently investors investing across these countries do not face exchange rate risk and have access to a wider range of bonds. However, the common monetary policy is not the only thing these markets share: the

[3]Finding differing credit risk premium in nominal and inflation-indexed yields of the same issuer would comply with the notion of partial or selective default, a possibility allowed by numerous macroeconomic models, like in Barro (2006) and Bolton and Jeanne (2009).

institutional features, market conventions, even the market (micro)structure of these products are similar across these three countries.

Although in the past issuance via syndication was a common practice, nowadays both nominal and inflation-linked bonds are issued via auctions of the corresponding Treasury agencies: the German Finanzagentur GMbH, Treasury Department of the Ministry of Finance (Dipartimento del Tresoro) in Italy and the Agence France Tresor. These auctions, identical to those in the US, are open to primary dealers, institutional investors who buy these assets. These institutions, typically either directly or through subsidiaries, participate in markets of all three countries. After the issuance and often multiple re-openings, these bonds are traded on the OTC secondary market, which in Europe consists of a handful of trading platforms. Most of the platforms trade all securities, however, there is some degree of specialization among them.

Nevertheless, it is not only the way these securities are traded that is similar in this cross-country sample. These products also have the same market conventions. This is especially interesting for inflation linked bonds in this study. The inflation-linked bond markets of these three countries are among the largest inflation-linked market segments of the world (Fleckenstein et al., 2014), their total value ($450 million) is half of the corresponding US segment. An interesting feature of the ILBs in our sample is that they are linked to one price index, to the Harmonized Index of Consumer Prices, henceforth HICP. This index is the weighted average of inflation of Eurozone countries and is published by the European Central Bank on a monthly basis. Moreover, the same deflation option applies to them: the principal payment of these bonds is protected when deflation occurs.

These countries started to issue ILBs in the past two decades. First, France issued inflation-linked bonds in 1998, a year after the US, but those were indexed to the French Consumer Price Index. Later, in 2001 they added HICP-linked bonds to their range of products. These bonds were especially popular among institutional investors across the Eurozone, as they were the first to compensate for Eurozone inflation with moderate sovereign risk at that time. Since 2003, the Ministry of Economy and Finance in Italy has also been issuing HICP indexed bonds. Today, they have the largest outstanding inflation-linked debt in the Eurozone. And at last, the German Finanzagentur has also issued its first ILB in 2006, and Germany was the first to issue an ILB after the financial crisis in 2008.

In conclusion, these three countries constitute a unique cross-section to study the relative pricing of real and nominal bonds. The fore mentioned features are the same across all bonds in the sample, therefore it is unlikely that any convenience yield would arise due to differences in trading or market conventions. Moreover, despite sharing the same currency and monetary policy, these countries still have different fiscal behavior, which results in diverse risk exposures of these debt securities.

1.2.2 Identification of liquidity and credit risk effects

The simplest way to quantify liquidity and credit effects in bond yields is to look at the individual asset markets in each country and estimate models with the corresponding risk factors separately. This can be applied to both indexed and nominal bonds. Practically this means that we would estimate a separate model for each bond segment: altogether six models in this cross-country sample. The clear advantage of this method is the direct comparability to results from the US Treasury market, as in Pflueger and Viceira (2015) or in Driessen et al. (2014). However, the major shortcoming is that not only one has to impose a lot of structure and assumptions on the estimation, but also that we cannot efficiently measure the relative riskiness of real and nominal bonds by only comparing risk exposures and price of a certain risk among different segments. Moreover, estimation might be infeasible due to insufficient data in segments with short time series and small cross-sections, such as the German ILB segment.

The methodological innovation in this paper is to directly estimate relative risk exposures of nominal and inflation-linked bonds and prices of their differential risk exposures. For the latter we model the holding period return of a single asset as a combination of market, liquidity and sovereign credit risk exposures. The next section explains how this pricing relationship applies to breakeven rates, whereas Appendix 1.A links it to the trading rule in Fleckenstein et al. (2014). An implicit assumption of the analysis is that Eurozone bond markets are integrated, which in light of the monetary union, common currency and other features of these markets is fairly reasonable. However, assuming integration not only has interesting economic implications but also a crucial role in the aggregation: it helps to restrict the number of parameters in the estimation and allows for the identification of the model in the cross-section of breakeven rates. Therefore, we define the market return as the equally weighted average return[4] of all bonds: inflation-linked and nominal bonds.

Liquidity is a multifaceted concept; both the level of asset and market liquidity (Amihud and Mendelson, 1986; Amihud, 2002; Bekaert et al., 2007) and liquidity risk (Pastor and Stambaugh, 2003; Acharya and Pedersen, 2005; Schwarz, 2015; Driessen et al., 2014) are likely to be priced. Moreover,Driessen et al. (2014) show that the importance of the level and risk aspects of liquidity differ across TIPS, nominal Treasury and inflation swap

[4]Equal weighting over-represents the smaller ILB segment, potentially exposing the market factor to liquidity and credit risks, which would weaken our results. As a robustness check, we will introduce a value-weighted market factor.

markets in the US. Therefore, we include both features: the level of liquidity of an asset is proxied by bond characteristics, such as age or size of an issue, whereas liquidity risk is captured by a liquidity factor. Moreover, we also control for sovereign credit risk. Most studies that examine credit risk look at the differences across countries (Arce et al., 2011; Beber et al., 2009; Ejsing et al., 2012) Despite that these differences are pronounced and highly economically significant around major credit events, such as the Euro crisis, looking at within country dissimilarities can be equally interesting: a country could choose to default on certain types of obligations, but not or to a different extent on others. This selective default can manifest in delayed payments, restructuring or the refusal of any payments to groups of creditors.

A fairly recent historical example[5] described by Duffie et al. (2003), is Russia defaulting on its ruble-denominated internal debt in 1998, whereas not on its eurobonds, shows that the occurrence of such event might not be unlikely. Moreover, eurobonds are similar to inflation-linked debt in nature, as the exchange rate and inflation risks both introduce uncertainty concerning the future payments that the issuer has to deliver. Inspired by this anecdotal evidence, we examine whether selective default risk is priced after controlling for liquidity and market risk exposures. To do so, we propose to describe expected returns in each market segment of the three countries with the following relationship, where all risk factors measure Eurozone-wide risks:

$$R_{i,t} - R_{\mathrm{f},t} = \alpha_i + \beta_{\mathrm{MKT},i} \left(R_{\mathrm{MKT},t} - R_{\mathrm{f},t} \right) + \beta_{\mathrm{LIQ},i} \eta_t + \beta_{\mathrm{CR},i} \theta_t + \varepsilon_{i,t}, \tag{1.1}$$

$$\mathbb{E} \left(R_{i,t} - R_{\mathrm{f},t} \right) = \kappa \mathbb{E} \left(Liq_{i,t} \right) + \lambda_{\mathrm{MKT}} \beta_{\mathrm{MKT},i} + \lambda_{\mathrm{LIQ}} \beta_{\mathrm{LIQ},i} + \lambda_{CR} \beta_{\mathrm{CR},i}. \tag{1.2}$$

In the above equations $\beta_{\mathrm{MKT},i}$, $\beta_{\mathrm{LIQ},i}$ and $\beta_{\mathrm{CR},i}$ are exposures to market, liquidity and sovereign credit risk factors, respectively. η_t and θ_t are the liquidity and credit risk factors, and $\mathbb{E}\left(Liq_{i,t}\right)$ captures the level of liquidity, proxied by asset characteristics. λ_{MKT}, λ_{LIQ} and λ_{CR} are the market, liquidity and credit risk premia.

To directly test the proposition of selective default, one has to compare the prices of credit risk in the nominal and inflation-indexed bond markets. If these two prices were not equal, that would provide evidence that nominal and indexed bonds are exposed to credit risk to a different extent. The next section presents a more direct approach, with which selective default risk can be directly measured.

[5] More recent examples are Argentina defaulting on its dollar denominated debt from the IMF in 2002, or the default of Ecuador in 2008.

1.2.3 The spread on breakeven (SBEI) strategy

In order to directly measure risk exposures, we propose to estimate risk premia based on pairs of breakeven rates. Breakeven rate or breakeven inflation is the yield difference between a nominal and real yields of bonds with similar maturities and credit quality. This yield spread is often thought of as a proxy for inflation expectations (e.g. Ciccarelli and Garcia (2009)), however it contains convexity and compounding effects (Kerkhof, 2005), inflation risk premia (Gürkaynak et al., 2010; Grishchenko and Huang, 2013) and other risk premia, such as compensation for liquidity risk (D'Amico et al., 2010; Pflueger and Viceira, 2015). Looking at the breakeven rate in one country is informative, nevertheless, taking the difference between pairs of breakeven rates across countries allows me directly identify relative risk premia in the underlying bonds.

Most studies that analyze the breakeven rate rely on the difference between two smooth zero coupon curves. As opposed to this, we choose to focus on pairs of bonds with the smallest possible maturity mismatch[6] between potential pairs across countries. We do this because on the one hand this allows me to use observable yields, therefore incorporate market information; on the other hand, we have to impose less assumption on the data as we are not fitting yield curves. Additionally, fitting the real curve could be challenging at the country level due to insufficient number of cross-sectional data points. Also, pairs of breakeven rates are practically bond portfolios with long and short positions. Consequently, we can show that the asset level models from the previous section can be aggregated to the portfolio level, additionally, as a result similar pricing relationships arise. Appendix 1.A shows, that one could derive the same model and pricing relations based on the trading rule of Fleckenstein et al. (2014).[7]

Differencing breakeven rates eliminates common components, such as 1) the compounding and a large part convexity effects that arise due to inflation; 2) the effect of inflation expectations and inflation risk premia; 3) any other factors that are the same across the three Eurozone countries, such as the effect of monetary policy or market or interest rate risks. The residual that the differencing does not take out is the exposure to risks that do not affect nominal and inflation-linked bonds equally within and across countries. While testing relative risk exposures, we also examine the assumption underlying the literature that studies relative liquidity of inflation-linked and nominal bonds that these bonds have identical sovereign credit risk exposures. Furthermore, this differencing based strategy

[6] In matching maturities, we follow Fleckenstein et al. (2014), however, one could also develop dynamic strategies based on matched duration or minimizing the convexity gap between nominal and inflation-linked bonds.

[7] In matching maturities, we follow Fleckenstein et al. (2014), however, one could also develop dynamic strategies based on matched duration or minimizing the convexity gap between nominal and inflation-linked bonds.

not only allows me to study liquidity and credit risk premia in sovereign bond prices alongside with excluding the above alternative explanations, but it also proposes a more stringent test of the relative pricing of inflation indexed and nominal bonds.

I propose that after excluding a battery of common components by the differencing, the residual yields are most likely driven by liquidity and sovereign risk differences – both within and across countries. Note that we can only identify significant credit premium in either one of these cases: 1) if the loadings of such premium differ across nominal and indexed bonds, or 2) if the price of credit risk differs between ILBs and nominal bonds. This latter possibility could arise due to selective default risk premium. Yet, similarly to liquidity risk, we focus on the effect triggered by Eurozone-wide credit shocks, due to the fore mentioned identification restrictions. For notational simplicity, we show how to quantify differential liquidity and credit effects from a maturity-matched German and Italian bond pair:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] &= \mathbb{E}\left[\left(R_{\mathrm{nom},t}^{\mathrm{G}} - R_{\mathrm{rep},t}^{\mathrm{G}}\right) - \left(R_{\mathrm{nom},t}^{\mathrm{IT}} - R_{\mathrm{rep},t}^{\mathrm{IT}}\right)\right] \\ &\approx \left(y_{\mathrm{nom},t}^{\mathrm{G}} - y_{\mathrm{ILB},t}^{\mathrm{G}}\right) - \left(y_{\mathrm{nom},t}^{\mathrm{IT}} - y_{\mathrm{ILB},t}^{\mathrm{IT}}\right). \end{aligned} \tag{1.3}$$

where $R_t^{country}$ stands for the return on the country-level breakeven or bond portfolio. This return, can be proxied by the yield difference of nominal and inflation-linked bonds, following Campello et al. (2008), who treat yield-to-maturity of a bond as a forward-looking expected return proxy. Then if we apply the pricing relation of Equation 1.2 to all four bonds underlying the SBEI strategy, we get the following:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] =& \kappa_i^{\mathrm{G,nom}} Liq_{i,t}^{\mathrm{G,nom}} + \beta_{\mathrm{MKT},i}^{\mathrm{G,nom}}\lambda_{\mathrm{MKT},t}^{\mathrm{G,nom}} + \beta_{\mathrm{LIQ},i}^{\mathrm{G,nom}}\lambda_{\mathrm{LIQ},t}^{\mathrm{G,nom}} + \beta_{\mathrm{CR},i}^{\mathrm{G,nom}}\lambda_{\mathrm{CR},t}^{\mathrm{G,nom}} \\ & - \kappa_i^{\mathrm{G,ILB}} Liq_{i,t}^{\mathrm{G,ILB}} + \beta_{\mathrm{MKT},i}^{\mathrm{G,ILB}}\lambda_{\mathrm{MKT},t}^{\mathrm{G,ILB}} + \beta_{\mathrm{LIQ},i}^{\mathrm{G,ILB}}\lambda_{\mathrm{LIQ},t}^{\mathrm{G,ILB}} + \beta_{\mathrm{CR},i}^{\mathrm{G,ILB}}\lambda_{\mathrm{CR},t}^{\mathrm{G,ILB}} \\ & - \kappa_i^{\mathrm{IT,nom}} Liq_{i,t}^{\mathrm{IT,nom}} + \beta_{\mathrm{MKT},i}^{\mathrm{IT,nom}}\lambda_{\mathrm{MKT},t}^{\mathrm{IT,nom}} + \beta_{\mathrm{LIQ},i}^{\mathrm{IT,nom}}\lambda_{\mathrm{LIQ},t}^{\mathrm{IT,nom}} + \beta_{\mathrm{CR},i}^{\mathrm{IT,nom}}\lambda_{\mathrm{CR},t}^{\mathrm{IT,nom}} \\ & + \kappa_i^{\mathrm{IT,ILB}} Liq_{i,t}^{\mathrm{IT,ILB}} + \beta_{\mathrm{MKT},i}^{\mathrm{IT,ILB}}\lambda_{\mathrm{MKT},t}^{\mathrm{IT,ILB}} + \beta_{\mathrm{LIQ},i}^{\mathrm{IT,ILB}}\lambda_{\mathrm{LIQ},t}^{\mathrm{IT,ILB}} + \beta_{\mathrm{CR},i}^{\mathrm{IT,ILB}}\lambda_{\mathrm{CR},t}^{\mathrm{IT,ILB}}. \end{aligned} \tag{1.4}$$

Note that the above betas are asset level risk exposures and can be estimated at the asset level. To get these betas we regress excess individual bond returns on the three risk factors of Equation 1.1. Then we impose that the above yields can be described as a combination of factor exposures and their respective premia, as in Equation 1.2. This is similar to the second stage of the Fama-MacBeth approach. In what follows, we apply these assumptions and we conjecture that the level of liquidity has the same coefficients across all assets in the strategy.[8] Consequently, the equation simplifies to:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] =& \kappa_i \left(Liq_{i,t}^{\mathrm{G,nom}} - Liq_{i,t}^{\mathrm{G,ILB}} - Liq_{i,t}^{\mathrm{IT,nom}} + Liq_{i,t}^{\mathrm{IT,ILB}}\right) \\ &+ \left(\beta_{\mathrm{MKT},i}^{\mathrm{G,nom}} \lambda_{\mathrm{MKT},t}^{\mathrm{G,nom}} - \beta_{\mathrm{MKT},i}^{\mathrm{G,ILB}} \lambda_{\mathrm{MKT},t}^{\mathrm{G,ILB}} - \beta_{\mathrm{MKT},i}^{\mathrm{IT,nom}} \lambda_{\mathrm{MKT},t}^{\mathrm{IT,nom}} + \beta_{\mathrm{MKT},i}^{\mathrm{IT,ILB}} \lambda_{\mathrm{MKT},t}^{\mathrm{IT,ILB}}\right) \\ &+ \left(\beta_{\mathrm{LIQ},i}^{\mathrm{G,nom}} \lambda_{\mathrm{LIQ},t}^{\mathrm{G,nom}} - \beta_{\mathrm{LIQ},i}^{\mathrm{G,ILB}} \lambda_{\mathrm{LIQ},t}^{\mathrm{G,ILB}} - \beta_{\mathrm{LIQ},i}^{\mathrm{IT,nom}} \lambda_{\mathrm{LIQ},t}^{\mathrm{IT,nom}} + \beta_{\mathrm{LIQ},i}^{\mathrm{IT,ILB}} \lambda_{\mathrm{LIQ},t}^{\mathrm{IT,ILB}}\right) \\ &+ \left(\beta_{\mathrm{CR},i}^{\mathrm{G,nom}} \lambda_{\mathrm{CR},t}^{\mathrm{G,nom}} - \beta_{\mathrm{CR},i}^{\mathrm{G,ILB}} \lambda_{\mathrm{CR},t}^{\mathrm{G,ILB}} - \beta_{\mathrm{CR},i}^{\mathrm{IT,nom}} \lambda_{\mathrm{CR},t}^{\mathrm{IT,nom}} + \beta_{\mathrm{CR},i}^{\mathrm{IT,ILB}} \lambda_{\mathrm{CR},t}^{\mathrm{IT,ILB}}\right). \end{aligned} \tag{1.5}$$

Nevertheless, if we wanted to quantify the respective risk premia from Equation 1.5, we would have to estimate nine of them. Given the limited number of maturity-matched basis pairs in the cross-section, we need to restrict the number of parameters to identify the regressions. Therefore, we focus our attention on cases where all risks are integrated at the Eurozone level. We do this by restricting the price of market, liquidity and credit risks to be equal across the four market segments in the SBEI pairs. Economically this means that liquidity and credit risk exposures are consistently priced in the cross-section of Germany, France and Italy. Ultimately we get the following relationship:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] =& \kappa_i \left(Liq_{i,t}^{\mathrm{G,nom}} - Liq_{i,t}^{\mathrm{G,ILB}} - Liq_{i,t}^{\mathrm{IT,nom}} + Liq_{i,t}^{\mathrm{IT,ILB}}\right) \\ &+ \lambda_{\mathrm{MKT},t}\left(\beta_{\mathrm{MKT},i}^{\mathrm{G,nom}} - \beta_{\mathrm{MKT},i}^{\mathrm{G,ILB}} - \beta_{\mathrm{MKT},i}^{\mathrm{IT,nom}} + \beta_{\mathrm{MKT},i}^{\mathrm{IT,ILB}}\right) \\ &+ \lambda_{\mathrm{LIQ},t}\left(\beta_{\mathrm{LIQ},i}^{\mathrm{G,nom}} - \beta_{\mathrm{LIQ},i}^{\mathrm{G,ILB}} - \beta_{\mathrm{LIQ},i}^{\mathrm{IT,nom}} + \beta_{\mathrm{LIQ},i}^{\mathrm{IT,ILB}}\right) \\ &+ \lambda_{\mathrm{CR},t}\left(\beta_{\mathrm{CR},i}^{\mathrm{G,nom}} - \beta_{\mathrm{CR},i}^{\mathrm{G,ILB}} - \beta_{\mathrm{CR},i}^{\mathrm{IT,nom}} + \beta_{\mathrm{CR},i}^{\mathrm{IT,ILB}}\right). \end{aligned} \tag{1.6}$$

And finally, by relabeling the portfolio of betas and liquidity characteristics as net effects, Equation 1.7 becomes a multifactor model inspired by Fama and French (1993) and Acharya and Pedersen (2005)'s Liquidity CAPM[9]:

$$\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] = \kappa_i \left(Liq_{i,t}^{\mathrm{net}}\right) + \lambda_{\mathrm{MKT},t}\left(\beta_{\mathrm{MKT},i}^{\mathrm{net}}\right) + \lambda_{\mathrm{LIQ},t}\left(\beta_{\mathrm{LIQ},i}^{\mathrm{net}}\right) + \lambda_{\mathrm{CR},t}\left(\beta_{\mathrm{CR},i}^{\mathrm{net}}\right). \tag{1.7}$$

The next section presents the estimation and gives a detailed explanation on how these equations are applied to the data.

[8]One of the robustness tests relaxes this assumption by allowing kappa to depend on the size of the underlying bond segments, which takes into account the relative size differences of nominal and indexed bond segments.

[9]In their paper net beta refers to the sum of market and the three distinctive liquidity betas that they estimate. Otherwise they also estimate a multifactor model with systematic and liquidity risks.

1.3 Estimation strategy

This section presents the data and describes their various sources. It is followed by the presentation of the main variables: the different liquidity and credit risk measures, the risk factors and a note on how expected returns are proxied. Finally, we give a detailed description of the estimation of both the market segment-level and breakeven-based strategies.

1.3.1 The data

The data are coming from different sources. The daily mid-quotes of nominal and inflation-linked bond prices are from Bloomberg, alongside with information on individual bond issues, such as issue and redemption dates, amount issued and coupon rates. The sample contains all available HICP-linked inflation indexed issues from the three countries: 5 from Germany, 9 French and 13 Italian ILBs. We focus on these assets as in the euro-area both inflation swaps and many inflation-indexed bonds are linked to this harmonized price index, while both Italy and France issue index linkers that are indexed to local inflation indices. However, having the same price index is crucial for the identification strategy. Alongside with inflation-linked debt, the sample covers a wide range of nominal issues, approximately 50-60 bonds from each country. The maturity dates of these bonds typically range between 2005 and 2055 and daily closing prices are adjusted by accrued interest following the respective market conventions. We collect data for the period between July 2004 and February 2014.

To capture the price effect of liquidity and credit risks, we complement the above data with 5-year sovereign quanto CDS prices for the credit risk factor, next to additional controls, such as the VIX and its European equivalents, the EURIBOR and EONIA indexes from Bloomberg. In order to define liquidity measures, we obtain the 10-year KfW agency bond yields and that of the 10-year constant maturity German nominal bond index from Datastream. To construct the benchmark liquidity proxy, we get data on monthly aggregate primary dealer transaction volumes directly from the German Finanzagentur and the Italian Dipartimento del Tresoro.[10] These figures are based on reports submitted by primary dealers on all transactions with other such institutions or third parties. Then these numbers are aggregated across counterparties and over the month and are available for the nominal and indexed segments separately.

[10] In one of the robustness tests we also use volume data for French bonds, which was obtained from the Agence France Tresor.

In the Eurozone, German bonds are argued to be the safest, therefore we use the 6-month constant maturity German sovereign yield as the risk free rate in our sample. Unfortunately, there are no bills issued with maturities shorter than 6 months, thus by imposing the assumption of bills having a flat term structure, we use it as the proxy for the 1-month rate to match the implicit holding period of the regressions.

1.3.2 Main variables

Asset, market and expected returns

Bond returns are the ratio of consecutive prices corrected for coupon payments. Market wide returns are based on the implicit assumption of Eurozone integration, and are defined as the equally weighted average across all bonds in the three countries. Standard asset pricing tests are usually performed on realized excess returns. As opposed to this, we quantify the effect of liquidity from bond yields, following Campello et al. (2008); Bongaerts et al. (2011); Pflueger and Viceira (2015); Driessen et al. (2014). We do so because yields are more persistent and less noisy than realized return estimates of a short sample. Under a set of assumptions, bond yields can be treated as forward-looking expected return proxies. First assumption is that markets are frictionless and that the term structure of expected returns is flat. For nominal bonds this relationship holds under the condition that yields follow a random walk process. As for ILBs, we also propose that inflation is constant in expectation and it is independently and identically distributed with yields. Absent liquidity and credit effects, one could show that the swap rate equals the breakeven rate. That case it can be proxied by the difference of nominal and real yields, therefore with the difference between two random walk processes that also follows similar dynamics.

Liquidity and credit measures

To explore the effect of liquidity, we include both asset and market level liquidity measures in the analysis. As in Fleckenstein et al. (2014) and Driessen et al. (2014), we face a similar problem that the directly observable bid-ask spreads are indicative quotes, thus not reliable and the Roll measure also cannot be constructed as bond returns tend to be positively serially correlated. Moreover, due to data availability, constructing the same set of measures for the six bond segments and the swap market is also infeasible. Therefore we proxy bond liquidity by using issue characteristics, such as age or amount issued, following Houweling et al. (2005). The reasoning behind a bond's age capturing liquidity is simple: the more time passes since issuance, the more likely that a bond gets locked-up in buy-and-hold investors' portfolios. This decreases its liquidity, which suggests a positive relationship between illiquidity and age, whereas issued amount is negatively related to the latter: larger issues tend to be more liquid. We define age as the years passed since issuance, whereas we use the natural logarithm of the amounts issued.

I also construct market wide liquidity measures that serve as a basis for the risk factor construction. One such proxy is the ILLIQ measure of Amihud (2002).[11] I define the measure as the ratio of monthly absolute bond market returns over monthly aggregate trading volume, where the volume is aggregated across all dealers and all securities within their segment and is observable at the monthly frequency. The second measure that we incorporate in the analysis is the KfW spread, which like Schuster and Uhrig-Homburg (2013)and Schwarz (2015), we define as the yield difference between a German agency bond issued by the Kreditanstalt fur Wiederaufbau and the maturity-matched nominal government bond. In constructing this liquidity spread, we follow Longstaff (2004) who quantifies liquidity premium as the yield difference between two securities that have the same credit risk but differ in their respective liquidities. Nevertheless, another potential interpretation of this measure is that it captures breakup risk or selective default risk. If this was the case, then using this spread as a liquidity measure could capture part of the credit risk premium in prices, which would result in an underestimated credit premium.

To capture each country's credit risk, we collect quotes from 5-year quanto CDS contracts. We use the changes in levels of the spread to construct the credit risk factor. Appendix 1.B provides graphs of the time-series of the different ILLIQ measures, the KfW spread, swap market measures and the three CDS spreads. Next to the previous liquidity and credit proxies, we construct additional controls that are included in some of the robustness checks, such as yield volatility or a control for the slope of term structure of bonds. Yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is taken over the different maturities for a given month. This definition is the same across both swaps and bonds. For bonds we also include time-to-maturity, which is defined as the remaining years until maturity of a given issue. This variable controls for a maturity structure and incorporates the slope effect of the term structure of bonds.

[11] The ILLIQ measure used in this study is based on holding period returns, while Amihud (2002) uses price returns. However, this does not influence the main results of the analysis as the resulting measures and risk factors based on the two types of prices are virtually identical.

Liquidity and credit risk factors

In order to examine Eurozone integrated effects of liquidity and credit risks, we construct factors that incorporate the country-level measures, and take out their variation by using principal component analysis. The Eurozone-wide liquidity measure consists of the four ILLIQ measures from Italian and German markets and the KfW spread. All the above measures are formulated so that the factor loadings ensure they all capture illiquidity. Similarly, to get an integrated credit risk measure, we take the first principal component of the individual measures from the three countries. In both cases the first principal components capture the most part of the variation, and serve as input for the factor construction. We define the risk factors as the unexpected or surprise component of these persistent measures:

$$Factor_t = M_t - \mathbb{E}\left[M_{t-1}\right], \quad \text{where} \quad M_t = [\eta_t, \theta_t] \tag{1.8}$$

The above residual defines the risk factor: the difference between M and its expectation in the preceding period. To compute these innovations, we impose a first order autoregressive structure on the different principal components capturing both liquidity and the credit measures in the sample.

1.3.3 Estimation method

In this section we explain how liquidity and credit risks affect asset returns: how Equations 1.2 and 1.7 are applied to the data. For this, we first estimate bond level betas to measure risk exposures, then in the second step we aggregate these betas to measure the price of relative risk.

Bond betas

To estimate the relative risk exposures from the breakeven rates, we turn to the standard two-step procedure based on Fama and MacBeth (1973). Unlike in most asset pricing tests, we do not sort our assets into portfolios, as we are interested in their individual characteristics and this way we can also take advantage of their larger cross-sectional variation. However, this comes at the cost of having less precise beta estimates. Thus in the first stage we estimate the betas from bond level time-series regressions where we regress bond excess returns on the liquidity and credit risk factors, η_t and θ_t, respectively:

$$\begin{aligned} &R_{i,t} - R_{\mathrm{f},t} = \alpha_i + \beta_{\mathrm{MKT},i}\left(R_{\mathrm{MKT},t} - R_{\mathrm{f},t}\right) + \beta_{\mathrm{LIQ},i}\eta_t + \beta_{\mathrm{credit},i}\theta_t + \varepsilon_{i,t}, \\ &\text{for} \quad t = 1, 2, \ldots, T \quad \text{for each } i\text{, market and country in the sample.} \end{aligned} \tag{1.9}$$

Equation 1.8 showed that the risk factors are residuals from autoregressive regressions. We estimate betas and risk loadings for each nominal and inflation-linked bond in our sample. We restrict our attention to integrated risk premia, where the market, liquidity and credit betas capture a common, Eurozone-wide risk exposure to the underlying factors. Given the liquidity and credit measures, we are able to measure an asset's covariation with the integrated market liquidity and credit risk. The former captures the same facet of liquidity risk as Pastor and Stambaugh (2003), whereas the credit beta proxies the exposure to the average sovereign credit risk in the Eurozone. These covariances suggest that market liquidity and credit risks affect required returns positively, such that the more illiquid or credit risky a bond is, the higher returns investors expect, which decreases the asset's price.

Breakeven betas and the price of differential risk exposures

Given the limited number of available breakeven pairs, identification and estimation of the betas and risk factors is nontrivial. If we wanted to conduct the usual Fama-MacBeth procedure, in the first stage we would need to regress the spread on breakeven rates on country-level market and illiquidity and credit risk factors from Germany and Italy, η_T^{G}, η_t^{IT}, θ_t^{G} and θ_t^{IT}, respectively to get beta estimates. However, there is no need to do this, as in the previous step we have already estimated the respective risk exposures based on Equation 1.9. Moreover, we are only interested in loadings on Eurozone risks – the ones that are common and likely to play an important role in both countries in the strategy. Therefore, we calculate the net betas from the bond level regressions the following way:

$$\begin{aligned} \hat{\beta}^{\mathrm{net}}_{\mathrm{G-IT,MKT},i} &= \hat{\beta}^{\mathrm{G,nom}}_{\text{EU-MKT},i} - \hat{\beta}^{\mathrm{G,ILB}}_{\text{EU-MKT},i} + \hat{\beta}^{\mathrm{IT,nom}}_{\text{EU-MKT},i} - \hat{\beta}^{\mathrm{IT,ILB}}_{\text{EU-MKT},i}, \\ \hat{\beta}^{\mathrm{net}}_{\mathrm{G-IT,LIQ},i} &= \hat{\beta}^{\mathrm{G,nom}}_{\text{EU-LIQ},i} - \hat{\beta}^{\mathrm{G,ILB}}_{\text{EU-LIQ},i} + \hat{\beta}^{\mathrm{IT,nom}}_{\text{EU-LIQ},i} - \hat{\beta}^{\mathrm{IT,ILB}}_{\text{EU-LIQ},i}, \\ \hat{\beta}^{\mathrm{net}}_{\mathrm{G-IT,CR},i} &= \hat{\beta}^{\mathrm{G,nom}}_{\text{EU-CR},i} - \hat{\beta}^{\mathrm{G,ILB}}_{\text{EU-CR},i} + \hat{\beta}^{\mathrm{IT,nom}}_{\text{EU-CR},i} - \hat{\beta}^{\mathrm{IT,ILB}}_{\text{EU-CR},i}. \end{aligned} \tag{1.10}$$

The net betas are portfolios of the respective risk exposures of the two nominal and indexed bonds that comprise the SBEI series. In this portfolio, the sign of each bond is according to that of the position in the breakeven rate. We also control for asset level liquidity, as shown in Equation 1.7, which is constructed as a portfolio of asset level liquidity measures:

$$Liq_i^{\text{net}} = Liq_i^{\text{G,nom}} - Liq_i^{\text{G,ILB}} + Liq_i^{\text{IT,nom}} - Liq_i^{\text{IT,ILB}}. \tag{1.11}$$

This transformation is applied to the asset characteristics for which we have data on all four bonds in the strategy, such as amount issued, age or time-to-maturity. Then to run repeated OLS regressions, we substitute expected returns by their forward-looking empirical counterpart[12] , by the breakeven rates, and estimate the following regressions:

$$b_t^{\text{G}} - b_t^{\text{IT}} = \gamma_t^{\text{net}} + \kappa_t^{\text{net}} Liq_{i,t}^{\text{net}} + \lambda_{\text{MKT},t}\left(\hat{\beta}_{\text{MKT},i}^{\text{net}}\right) + \lambda_{\text{LIQ},t}\left(\hat{\beta}_{\text{LIQ},i}^{\text{net}}\right) + \lambda_{\text{CR},t}\left(\hat{\beta}_{\text{CR},i}^{\text{net}}\right) + \varepsilon_{i,t}^{\text{net}},$$
$$\text{for} \quad i = 1, 2, \ldots, N \quad \text{for each } t \text{ and basis pair in the sample,} \tag{1.12}$$

where $b_t^{country}$ is the yield difference between the respective ILB and nominal issues, thus the breakeven rate. Estimates from these repeated regressions are averages across time and errors include both a 12-month Newey-West correction and account for the averaging of the coefficients. Moreover, the resulting premium estimates are directly interpretable: they show how large a part of the yield difference is accounted for by the reward for all four bonds in the strategy being exposed to liquidity and credit risks. This is a direct measure of partial or selective default risk premium.

1.4 Empirical results

This section presents the results of this study. First, we show the descriptive statistics of the main variables, then proceed with reporting the estimated betas and the net or portfolio betas. We also discuss the time-series properties of the factors. Then we proceed, with the analyses of the relative pricing of nominal and indexed bonds. These are based on the direct approach following Equation 1.12. At last, we present various robustness checks, such as pooled OLS regressions, convexity calculation and trading volume weighted liquidity level estimation. We conclude this section with a discussion that touches upon the size of the credit effect, CDS liquidity, and macro implications and mechanisms behind the results.

[12] Appendix 1.A explains why this substitution is conceptually feasible.

1.4.1 Descriptive statistics, betas and factors

Table 1.1 contains descriptive statistics of the main variables, whereas Table 1.2 provides an overview of the beta estimates for both the benchmark and segmented market cases. In Table 1.1, Panels A to C compare the different features of nominal and inflation-linked bond. The main variables are in line with expectations: in ILB markets the yields are lower and, on average, less volatile than nominal ones, where the German average yield is the lowest. German ILBs are the youngest as these bonds are only issued since 2006, whereas nominal bonds are older in all three markets. In Germany the average size of nominal issues is almost 30% larger than ILBs, whereas in Italy and France this difference is even larger, 50% and 100%, respectively. We also present the ILLIQ measure that shows the absolute euro change in price triggered by trading 1 million EUR. This price impact is the highest in the German ILB and the lowest in the German nominal segments. This observation verifies that German nominal bonds are highly liquid, especially in times of flight to liquidity. Inflation swaps have an average yield of 2.19% whereas the difference between average indexed and nominal yields is 67 basis points in Germany, and 164 and 83 basis points in France and Italy, respectively.

Figure 3.1 depicts the time evolution of both the country and Eurozone level illiquidity and credit factors. All series have their peaks at the financial and the euro crises, which is in line with anecdotal and previous empirical evidence. The country level liquidity factors differ slightly: in Germany it is constructed by taking the first principal component of the KfW spread, and the ILLIQ and zero return measures from both the nominal and indexed segments. All of these measures have a positive loading in the first component, except for zero returns in the ILB market. However, this is not surprising in light of the segment being relatively young and there are a high number of zero return days in the months succeeding its introduction, but not later. For Italy, the principal component is based on the different ILLIQ and zero return measures, where the constituent measures show the same relation: all measures constituting the German and Italian factors are positively correlated to one another. Individual measures are depicted in Appendix 1.B. The three illiquidity factors from the bond markets follow similar dynamics and hence their correlations are sizeable: it is 0.41 between the German and Italian liquidity factor. The credit factors are based on the unexpected changes in the sovereign CDS series. These series tend to closely follow each other, as can be seen in Figure 1.B.5 and exhibit correlations above 90%. After taking the residuals from the respective autoregressive processes, the countrywide credit factors remain highly correlated: all coefficients are above 0.7.

Table 1.2 presents the distribution of beta coefficients across the six bond market segments and the net betas from Equation 1.13. The betas are estimated from asset-level time-series regressions of excess returns on the market, illiquidity and credit factors; under the assumption of either integrated or segment-level market factors.[13] Under these assumptions the market factor is the Eurozone or asset segment-wide equally weighted average return, respectively. In both cases we expect liquidity and credit betas to be negative on average, whereas the segmented market betas being close to one. On the one hand, this is not what we find in the data in all cases. Market betas in all segments are different from one and often negative. This is due to the non-homogeneous and imbalanced nature of the market factor, whose composition changes whenever a new issue enters or an old one reaching maturity leaves the sample. On the other hand, there is a pattern in nominal integrated market betas that is consistent with flight-to-quality: Italian nominal yields increase whenever European systematic risk rises, French bonds show only a slight effect, whereas the negative beta of the German nominal sector suggests that investor find safe haven in these assets. The other irregularity of the betas is that not all liquidity and credit betas are negative on average. In German and Italian markets this seems less of a problem, unlike in France, where we cannot construct a segment-specific French illiquidity factor to measure the respective beta. Instead we substitute the missing information with the integrated, Eurozone-level liquidity factor.

1.4.2 Net betas

Net betas can be found in Panel D of Table 1.2, as well as they are depicted in Figure 1.2. Net betas are a portfolio of nominal and ILB betas that constitute the spread on breakeven strategy. There are twenty such strategy pairs in the sample that have at least 12 monthly observations. Panel D shows that the average net beta is negative in all three cases, in addition, Figure 1.2 is also in line with this observation. Economically speaking, if the liquidity and credit risk exposure were the same among nominal and inflation-linked bonds, net betas would line up at the zero. Therefore, finding values other than zero suggests that exposures differ among the two bonds, moreover, this difference is also not consistent or the same across the two countries. Moreover, the sign of these betas also suggest which of the underlying four bonds drives the result, this can be derived from the sizes and signs of the individual bond betas.

Liquidity net betas can be found in a narrow range around zero, while credit net betas in the mid-panel are more dispersed and larger – often even by two orders of magnitude. The

[13] I also looked at median betas and observation-weighted average betas. This latter method would tilt the average towards more precise estimates, the ones that are based on more monthly observations. However, the resulting betas are not significantly different from the ones that are presented in the table: they are similar in magnitude and have identical signs.

negative liquidity betas suggest that ILBs are less liquid than nominal bonds, a finding in line with previous literature. However, one of the most surprising finding of this paper is that we find that credit risk exposures of bond within a country, issued by the same issuer, are also large enough to survive the double differencing. The sign of the credit betas also provides suggestive evidence of which bond is driving this relationship: ILBs are more exposed to sovereign risk, whereas there is also a natural ordering across the countries in terms of their riskiness: Germany is the safest from a sovereign perspective, Italy is the least creditworthy, whereas there is mixed evidence for France. Finally, the market net betas in the lower panel are the largest in size and dispersion, despite that one would expect such exposures to be zero. These loadings are a clear proof that the breakeven spread is exposed to integrated non-diversifiable Eurozone risk, similarly to the finding of Monfort and Renne (2014). This is due to the integrated market factor capturing some aspects of liquidity and credit risks in the euro area, which are apparently relevant in the pricing of the markets under scrutiny.

The beta estimates reflect how difficult it is to disentangle the effect of liquidity and credit risk, two concepts that are highly correlated and intertwined, especially in distressed periods. There have been many papers trying to separate them, and our approach is the best attempt to explore such a highly relevant question in this recently available cross-section of indexed and nominal euro area sovereign bond data.[14]

1.4.3 The relative pricing of indexed and nominal bonds

There is evidence from the US Treasury markets that both the level (Krishnamurthy, 2002; Goyenko et al., 2011; Fleckenstein et al., 2014; Pflueger and Viceira, 2015; Driessen et al., 2014) and risk (Driessen et al., 2014) aspects of liquidity are priced, whereas empirical findings from the Eurozone are restricted to nominal bonds (Darbha and Dufour, 2014; Pelizzon et al., minga; Schwarz, 2015). As opposed to this, to our best knowledge this is the first study to present empirical evidence for selective default risk premium in the relative pricing of nominal and inflation-linked bonds. Next to this, the main contribution of the paper is coming from the identification strategy that helps better understanding the relative pricing of inflation-linked and nominal bonds and to set up clean asset pricing tests in a difference-in-differences setting, ensuring that the analysis is the least contaminated by confounding effects.

Unlike the majority of the literature working with breakeven rates, we choose to focus on maturity-matched bonds we construct the matched pairs by minimizing the mismatch

[14] Beta estimates presented might be estimated with considerable errors due to the short and imbalanced sample in addition to estimating them from individual assets.

of the two bond maturities, as well as we try to match similar tenors to one another. Consequently, we have breakeven rates on 5 or 10 year or mixed maturities. Due to this heterogeneity in tenors, and the various contaminating effects that are eliminated by the differencing, studying these yield spreads are less informative then doing so based on their differences. Therefore, we focus the analysis on the SBEI series, for which we use the pool of 27 maturity matched bond pairs. The resulting series are depicted in Figure 1.3, where the different panels correspond to different country pairs. There are twenty SBEI series with at least 12 months of data available in the sample: 6 of these are taken between Germany and Italy, 5 pairs are formulated across German and French bond pairs and 9 pairs are among Italian and French breakeven rates. The descriptive statistics of these series are in Panel A of Table 1.4.

Panel A of Table 1.4 shows that the average breakeven spread is between -36 and 23 basis points across all the pairs. The lowest value the spread takes is -272 basis points, for Italian-French matched basis pair, whereas its peak is 401 basis points for a pair from Germany and Italy. The latter series are the most volatile. The economic interpretation of the German-Italian example based on the average series is that if an investor were engaged in German and Italian ILB and nominal positions according to Equation 1.3, her average return would amount to 23 yield basis points. However, the panel also shows that holding the underlying positions results in volatile returns with large swings ranging between a loss of 66 basis points or gains in the order of 4% per annum. This suggests that even without taking transaction costs into account, there are potential sizeable losses arising if the strategy cannot be held until maturity. This is even more prominent in German-French and Italian-French bond pairs. However, discovering the limits to the tradability and risk-return characteristics of the breakeven spread is not the main focus of the paper. Instead, Table 1.4 contains the results estimated from monthly repeated cross-sectional regressions of the SBEI series on the illiquidity and credit factors, alongside with composite asset level liquidity proxies, based on Equation 1.10 and 1.12. Net betas are portfolios of betas estimated from the first stage of market segment level Fama-MacBeth regressions, as loadings on the Eurozone market, illiquidity and credit factors. The results correspond to the period between July 2004 and February 2014.

This implies that (1) the spread on breakeven is exposed to systematic risk; (2) the European interest rate or integrated market factor is likely to be capturing some aspects of liquidity and credit risk. This is additional proof of how important these risks are in the pricing of sovereign bonds. Nevertheless, this finding shows that liquidity differences between nominal bonds and ILBs is probably more a within than an across-country effect.

Table 1.3 focuses on the effect of liquidity and sovereign risk differences between nominal and inflation-linked sovereign bonds. Column 1 considers liquidity alone, which has a highly significant discount. This effect, evaluated at the mean net liquidity beta, explains 3.48 basis points of the SBEI series. However, this effect is not robust to the inclusion of differential credit risk or bond group-level liquidity measures. The most striking and key finding of the paper is that differential credit risk is priced in the cross-section of Eurozone SBEI series. This implies that investors probably perceive ILBs riskier from a sovereign risk perspective and see governments to be more likely to default on their inflation-indexed obligations than on nominal debt. This observation is in accordance with selective default events that occurred in Russia in 1998 with its ruble-debt, or with Argentina between 2003 and 2005 that defaulted on eurobonds. Moreover, to our knowledge this finding makes our paper the first to provide empirical evidence on such phenomenon.

Finding significant credit effects can partly be explained by the existence of a non-diversifiable euro-area credit risk as in Monfort and Renne (2014), and the constituent assets' exposure to such a factor. However, identifying any credit effects that does not cancel out within a certain country suggests that this is selective default risk. The inclusion of credit risk almost doubles the R-squared compared to when only illiquidity is considered. Credit risk carries a large discount, its market price is -126 basis points; and accounts for a sizeable yield difference of 41.6 basis points evaluated at the mean net credit beta. This effect is persistent and robust to the inclusion of liquidity level proxies, among which the relative age of bonds also matters; and to incorporating systematic risk by means of the market factor.

As unusual as pricing differential risk exposures might seem, the idea of comparing yields of securities with similar exposures to certain risks is not new in the literature. Longstaff (2004) compares yields of US Treasuries to those of bonds issued by the Refcorp (Resolution Funding Corporation), whereas Schwarz (2015) examines yield differences of German federal government bonds and bonds issued by KFW, a government owned development bank. The key feature of theses agency bonds is that they have explicit government guarantees, and consequently the same credit risk as government bonds. However, the liquidity of government bonds is substantially higher and thus the yield difference measures general market liquidity conditions. What we do in this paper is similar but goes the other way around: while controlling for liquidity on both the nominal and inflation-linked bond markets the same way, we show that the remaining yield difference is attributed to sovereign risk. This idea is also consistent with the alternative interpretation of the Refcorp and KfW spreads - some say that these yield differentials, rather than capturing liquidity, can also be interpreted as breakup or selective default risk measures.

Figure 1.4 depicts the percentage yield risk premium due to relative illiquidity and credit risk in all available breakeven spreads. It is calculated as the product of the cross-sectionally estimated risk premia and the respective net betas. The upper panel shows the illiquidity premium estimated from Column 1 in the respective pairs, whereas the lower graph depicts both liquidity and credit premia based on Column 2 of Table 1.3. There are two noteworthy observations: 1) illiquidity is not robust to the inclusion of credit risk, as both the magnitudes and sign of SBEI-level effects change across the specifications; 2) the magnitude of the yield difference that is explained by credit risk is tenfold compared to that of illiquidity, and in certain cases it is up to 1.5%. The discussion at the end of this section reconciles the potential drivers of the size of the credit effect.

The final step of examining the SBEI series is to evaluate the effect of risk adjustments. This shows how would the series change if we took out the estimated liquidity and credit premia, as in Table 1.4 and Figure 1.5. Table 1.4 presents descriptive statistics of the average unadjusted series, in Panel A; and the liquidity, and liquidity and credit risk-adjusted series in Panel B and C, respectively. The idea underlying the adjustment is based on Equation 1.15, where we proxy the SBEI series by the portfolio of constituent bonds. To calculate the adjustment for each bond quadruplet, we sum up the asset level risk premia and deduct this sum from the spread. We apply this raw correction to all bond pairs in the sample and recalculate the average of the adjusted spread series. Panel B presents the specification when only illiquidity adjustment is applied. In all three country pairs we find that by taking out the estimated (non time-varying) risk premia, the spread shrinks considerably. Panel C shows the case when the adjustment is based on the sum of liquidity and credit effects. As opposed to liquidity, taking both liquidity and credit risks out deepens the spread. Also the extreme values show that the series are shifted downwards, further from their equilibrium level of zero.[15]

Figure 1.5 gives a visual representation of the moments in the table. The top panel shows the breakeven pairs matched between German and Italian bonds; the middle panel refers to those from Germany and France; whereas the bottom panel depicts French-Italian pairs. In all three panels the solid line refers to the average series, the dashed one to liquidity, whereas the dotted line to the composite adjustment. The average series presented are also imbalanced: their composition over which they are defined varies over time.

The upper panel shows that the German and Italian spreads have their peaks at times most likely coinciding with ECB intervention. Applying liquidity adjustment to this series does not have a large effect. In general, the figure suggests that the corresponding liquidity

[15]Note that this adjustment does not allow me to incorporate time variation in risk premia; therefore, we can only induce parallel shift in the series. Incorporating time variation would require splitting the sample or using rolling window estimation that are not feasible due to the empirical issues explained before.

corrections are not too large and probably average out. On the other hand, applying credit adjustment shifts the average further from zero: it deepens the yield difference if the respective bonds in the strategy. This is not surprising, as these pairs contain the smallest and the largest credit risk estimates and by definition of the portfolio formation, the sum of these effects is expected to be negative. The middle section demonstrates the average of the five Germany and France-based series, for which liquidity shifts the average upwards, closer to zero. This helps to explain the price difference by illiquidity differences, but taking credit risk into account moves the opposite direction. This result is in line with the selective default story, where investors perceive ILBs riskier than their nominal counterparts. And finally, the lower panel shows how the changing composition of the average series influences the results: in the first period where the two adjustments exhibit different dynamics, one matched bond pair is available. Once other pairs enter the sample, the effect of credit risk goes to the same direction as that of the liquidity correction: taking them out of the spread deepens the spread by pushing it further away from zero.

1.4.4 Robustness tests and discussion

This section presents robustness tests. Unreported analyses include models with different illiquidity and credit measures and additional controls. The effects of liquidity are stronger when a funding liquidity proxy, the OIS spread, is included, whereas defining the credit factor based on changes not levels of the CDS spread virtually produces the same results. Age and time-to-maturity are also considered as proxies for asset level liquidity, though results are mixed. Other tests of the benchmark specification of the SBEI cross-sectional regressions are the following: 1) Quantifying the size of convexity effects in BEI and SBEI series; 2) Easing the assumption that level liquidity effects have to be the same across all four underlying bonds; 3) Pooled OLS to partly overcome the statistical difficulties of the small number of breakeven spreads. This latter method does not allow for direct examination of relative prices but helps to convince the reader that these risks are important on the pricing of bonds in the sample. And finally, this sections concludes with some discussion regarding selective default, alternative interpretations of the results and the size of the credit effect.

Convexity in BEI and SBEI series

Kerkhof (2005) points out that part of the breakeven inflation captures compounding and convexity differences between nominal and inflation linked bond issues. Convexity is the second derivative or the curvature measure of the price function and it measures how the duration changes as interest rates (yields) change. As such, convexity is likely to play a role in the analysis presented. However, the definition of convexity is somewhat different across nominal and inflation-linked bonds: they are calculated with respect to nominal and real rates, respectively. There are different market conventions to determine the convexity of ILBs: 1) practitioners either base convexity on real yields and consider the pricing of ILBs in a world without inflation risk; or 2) they use an older method developed for UK Gilts, in which inflation assumptions are used to discount future cash flows of an indexed bond. In our analysis we use this latter method, in which we assume that the term structure of inflation is flat and constant at the level of 1.75%.[16] Then we calculate the convexity of a "quasi-nominal" bond that pays coupons and principal adjusted for inflation, and its sensitivity is evaluated against the real yield.

As the cornerstone of our analysis is the breakeven rate, we examine the difference in convexities within a country between nominal and inflation-linked bonds, as well as how large this effect is across pairs of bond pairs of the SBEI series. To quantify the return or yield effect due to convexity in nominal and indexed bonds, we multiply their respective convexities by 1/2 and the variance of the change in yields similar to Campello et al. (2008). We aggregate the product of the convexity effects and yield variances by summing up these bond level effects according to the bond positions in the SBEI quadruplets. To control for different levels of yield volatility in the nominal and indexed segment, we assume here for simplicity that the variance of nominal yield changes equals to 1%, while for ILBs we assume a value of 0.5%. The resulting calculation suggests that the yield variance differential of nominal and real bonds leads to an effect on expected return of 17.87 and 35 basis points for German-Italian pairs and German-French pairs, while it is 52.86 basis points for Italian-French breakeven pairs.

In light of the magnitude of the estimated differential liquidity and credit premiums, these convexity effects seem substantial, especially in a period where yields are particularly volatile. Even if convexities of indexed and nominal bonds are not all that different, accounting for the difference in yield volatility of these bonds can result in a sizeable convexity correction that could affect the analyses. Consequently, as an extension of the chapter, we are planning test the robustness of the cross-sectional relative prices of credit and liquidity risks to estimation based on convexity-adjusted SBEI yields.

[16] This number is the average Eurozone (HICP) inflation since 1999, reported by the European Central Bank and Eurostat.

Weighted kappa

Equation 1.7 shows that expected returns on the SBEI series are defined as a sum of differential credit and liquidity effects. This pricing equation assumes that the liquidity level effect of the four bond in the SBEI portfolios are identical. In this section we relax this assumption by adjusting this level effect by the size of the respective bond segments. We do so, as the cross-sectional coefficient can be interpreted as turnover, as in Amihud and Mendelson (1986). Turnover is typically defined as the ratio of traded volume of a given asset or market segment over the overall volume outstanding thereof. Since we do not have information on the overall size of the individual bond market segments, we substitute this by the monthly trading volumes by primary dealers. To account for the different size of nominal and real bond segments, we construct a measure that captures each segment's relative share of the overall traded volume of all six segments in the sample:

$$w_{i,t} = \frac{trading\ volume_{i,t}}{overall\ trading\ volume_t}, \tag{1.13}$$

where i corresponds to the bond segments and t to the month between January 2006 and December 2013. The assumption underlying this measure is that the three Eurozone issuers constitute all trading in the Eurozone sovereign markets and thus the shares can be treated as a rough proxy for the turnover of a given segment. The weights for nominal and inflation-linked bond segments can be seen in the upper and lower panels of Figure 1.6, respectively. Note that the percentages of the two panel together add up to a hundred percent. The figure reveals that the size and trading activity of nominal and inflation-linked segments differ substantially: nominal segments have more trading, with German Bunds being traded the most and indexed Bunds the least. Moreover, the relative weights are rather stable over time.

The weights of Figure 1.6 are used to calculate the net level of liquidity measure as a trading share weighted average of liquidity levels. To proxy for size of an issue, we use the average trading share as weight, while for age and time-to-maturity we use time-varying weights:

$$Liq_{i,t}^{\text{w,net}} = w_{\text{G,nom},t} Liq_{i,t}^{\text{G,nom}} - w_{\text{G,ILB},t} Liq_{i,t}^{\text{G,ILB}} + w_{\text{IT,nom},t} Liq_{i,t}^{\text{IT,nom}} - w_{\text{IT,ILB},t} Liq_{i,t}^{\text{IT,ILB}}. \tag{1.14}$$

Substituting the new, weighted level of liquidity measure to our pricing equation, Equation 1.7, the following relationship arises:

$$\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] = \kappa_t \left(Liq_{i,t}^{\mathrm{w,net}}\right) + \lambda_{\mathrm{MKT},t}\left(\beta_{\mathrm{MKT},i}^{\mathrm{net}}\right) + \lambda_{\mathrm{LIQ},t}\left(\beta_{\mathrm{LIQ},i}^{\mathrm{net}}\right) + \lambda_{\mathrm{CR},t}\left(\beta_{\mathrm{CR},i}^{\mathrm{net}}\right), \quad (1.15)$$

where $\kappa_{i,t}$ captures the level of liquidity premium, which takes into account the relative size difference between the four underlying bond markets. Table 1.5 reports the results for this robustness test, however since the volume data are only available between January 2006 and December 2013, these results correspond to a somewhat shorter period than thereof Table 1.3. Column 1 is directly comparable, despite the shorter sample period, the price of differential liquidity risk is lower and still insignificant, whilst that of the differential sovereign risk is very similar in magnitude with a sign consistent with expectations. Columns 2 to 4 contain trading volume-weighted liquidity level measures. The effect of differential credit risk is robust to the inclusion of the weighted size and age measures, however, its significance suffers once the weighted time-to-maturity is added to the model in Equation 1.8. Once the weighted time-to maturity measure is included, liquidity risk becomes highly statistically significant.

In general, we conclude that the results are rather robust to relaxing the assumption on the level of liquidity proxies for the SBEI series, where we assume that the size of the market segments are proportional to their respective trading volumes. Nevertheless, one should remember that these regressions shed light on the relative pricing of inflation-linked and nominal bonds, thus the estimated kappas cannot directly be interpreted as turnover, nor should they be used to back out the implied holding period of the underlying bond portfolios.

Pooled regressions

Disentangling liquidity and credit risks is a difficult task. Moreover, the sample of this study is relatively short and highly imbalanced: despite that it spans ten years, there is barely any assets that actually span the entire sample period. This is due to new bonds being issued and old ones reaching maturity on a regular cycle. The resulting instability in composition mostly affects the market factor, but is likely to also increase the standard errors of other beta estimates. In addition, some markets are rather young and only a few assets are traded, which makes it even more difficult to identify cross-sectional effects. One way to alleviate the problem of small samples, although this does not allow for direct comparison, is to pool all bonds together for the analysis. This helps to establish the relationship between bond yields and liquidity and credit risks, and hopefully also to convince the reader of the relevance of the differential pricing.

To alleviate the above problems, we run regressions that are pooled across all bond market segments and estimate euro area wide risk premia; results are presented in Table 1.6. The shows that Eurozone interest rate or market factor is positive and significant across all specifications. Including credit risk in the analysis strengthens the effect of illiquidity risk: they are priced and highly significant together; moreover, their effects are robust to the inclusion of any asset level liquidity proxy. Adding these two factors increases the R-squared considerably. The average betas of both credit and illiquidity are negative; therefore, the market wide illiquidity premium translates into a negative average effect. Despite this unexpected effect, we find positive impact on yields, which is in line with economic theory. Eurozone-wide credit risk has positive and stable average effect on yields, yet its impact is negative. These inconsistent signs among average and interquartile effects show that the betas are asymmetrically distributed.

These pooled regressions help to alleviate the burden of estimating market level models from short and highly imbalanced subsamples. On the one hand, the identification based on differencing of breakeven rates is very clean as the differencing takes out commonalities between market segments and countries, but the cross-section of assets is very small (20 pairs of pairs), which aggravates the estimation error in betas and cross-sectional risk premiums. Moreover, any estimated risk premiums are relative risk premiums, quantifying the differential exposures of nominal and indexed bonds. This aspect of relative pricing is the central question of the paper, nevertheless it limits our ability to further generalize results or to formulate absolute statements. On the other hand, pooling together all asset segments, we get to have more observations and any effects are absolute concepts, not in the context of relative risk exposures. However, the drawback is that we should control for all the effects that the differencing takes out in the other setting. To this extent, this gives rise to an omitted variable bias, which should be controlled for.

Discussion

There are two potential economic mechanisms that could give rise to differences in sovereign risk exposures of inflation-linked and nominal debt: selective default risk and correlation between high inflationary states of the econoour and default. However, these mechanisms are not trivial to disentangle in a purely empirical setting or by means of the current dataset. This discussion describes these two mechanisms, as well as it aims to add a note on the size of the estimated risk premia.

First, the section on the relative pricing of sovereign bonds refers to selective default risk premium. We define selective default as an event, in which a sovereign issuer chooses not to meet obligations on a class of bonds, while servicing her other debt. Practically selective default can range from rescheduling payments to non-payment, and it also serves as a category that rating agencies use to rate sovereign issuers.[17] Of course the question arises which type of bonds to default on. Sovereign defaults are rare events in the post-WW II era; moreover, there has not been a precedent for a sovereign issuer specifically defaulting on her inflation-linked debt. This is partly due to the novelty of inflation-linked products, however, the literature suggests that the choice might depend on reputation (Duffie et al., 2003; Gennaioli et al., 2014) or riskiness of debt.

In the context of this paper, both motivations could play a role: 1) inflation-linked debt is very similar to foreign-currency denominated debt in that it also bears an additional risk, namely inflation risk. Additionally, the specific sample period of the study suggests that selective default premium is identified from the increased sovereign risk of the euro area countries during the financial and subsequent euro crises. Consistent with this idea, although the likelihood of such an event is small, our analysis shows that investors could have attached positive probability to Eurozone issuers to decide to strategically default on their indexed debt. 2) the HICP-linked debt is primarily issued to foreign institutional investors. Both Italy and France have ILBs that are linked to local inflation, so issuing real bonds linked to the average Eurozone inflation seems to primarily target foreign investors or institutions that for some reason prefer to hedge against this "average" European inflation.

What scenarios could trigger partial default in the euro area? On the one hand, we can interpret the main result as risk premium in ILBs compensating for larger sovereign risk exposure or for selective default risk. This means that despite legal clauses like pari passu, holders of real bonds perceive their claims less senior than those of nominal issues. In this case the government or Treasury should communicate its intention to fulfill all obligations at equal terms, or ensure debt holders that a collective action clause applies: it allows the supermajority of bondholders (75%) to agree to a debt restructuring that is legally binding to all debt holders. Since 2013, as part of the European Stability Mechanism all bonds issued by Eurozone member countries and with maturities longer than one year have mandatory collective action clause. On the other hand, the finding can be stemming from inflation and default being highly correlated, which means that the inflation protection is the most valuable exactly in the state when the default occurs. This case the channel potentially driving the results is inflation, more specifically the difference between the HICP inflation (what ILBs in our sample pay) and the French, Italian or German inflation. In the past years there were periods when the HICP was above French or Italian inflation. In such a case, it is more expensive to pay back HICP-linked debt

[17] "An 'SD' rating is assigned when S&P Global Ratings believes that the obligor has selectively defaulted on a specific issue or class of obligations but it will continue to meet its payment obligations on other issues or classes of obligations in a timely manner." / Source: standardandpoors.com/

than its French RPI-linked counterpart. Although these countries do not have full control over their monetary policy, the decision to default on their debt is the issuers discretion. Thus there could be instances where inflation, an additional risk factor, together with financial difficulties of a country could trigger selective default on HICP-linked debt. This scenario is potentially aggravated if in light of the European debt crisis the ECBs anti-inflationary stance proved to be less credible to both the individual countries and holders of HICP-linked debt.[18]

Lastly, we would like to add a note on the relative importance of credit and liquidity in the relative pricing of indexed and nominal sovereign debt and the size of the credit effects. The many attempts in the literature show that disentangling liquidity and credit risk in sovereign bond yields is a non-trivial task. In the specific case of this analysis, there is a possibility that the credit factor is picking up some aspect of liquidity, flight-to-safety or maybe even a credit level effect. Then the presented method would overestimate the effect of credit risk at the cost of liquidity risk.

Despite the lack of empirical evidence on expected credit risk or credit level effect in these securities, we find it plausible to exist, similar to such effect corresponding to liquidity. Nevertheless, testing for it is non-trivial: there is no cross-sectionally variable measure of asset-level credit effect – ratings are given to countries not to their individual bond issues, and due to the lack of actual sovereign defaults in developed countries (Reinhart and Rogoff, 2009) estimating the probability of default or the loss given default requires a lot of assumptions. Additionally, although the EBA and ESRB conduct stress tests on average on a bi-annual basis, this is still a global, country-level proxy that does not allow for the estimation of the expected credit effect in the cross-section of bond segments or individual bonds issues. For this reason, since our focus is predominantly on the average credit effect measured by the product of the average net beta and the estimated cross-sectional risk premium for differential credit risk between nominal and inflation-linked bonds. However, we acknowledge that the size of our estimated effect can be party explained by expected credit risk, if most of it does not cancel out by the differencing across countries. If this effect persists, then finding a higher credit beta for ILBs suggests also higher expected credit level effect, which could prove the point that ILBs are implicitly less senior than nominal bonds of the same issuer.

The other potential bias could be due to the exposure of CDS contracts to liquidity risk, although our choice of instrument, the 5-year quanto CDS contract, ameliorates this problem. In the period between 2006 and 2014 there are two major trends in the

[18] A potential test of this underlying correlation between default risk and inflation would be to conduct a subsample analysis in which we restrict the sample to months where HICP is above French or Italian inflation.

CDS market: in the beginning of the sample period there is a large increase in the number of contracts and dollar volume traded in sovereign CDS. Especially around the onset of the Euro crisis, there is increased interest in these products, alongside with speculative trading, which have been said to destabilize sovereign debt markets of the GIIPS countries. As a reaction to this, in October 2011 the European Union introduced a rule to stabilize sovereign bond markets that banned naked CDS positions. This ban came into effect on November 1, 2012. Due to this regulatory change the trading volume of European CDS contracts drastically fell and the market shrunk to a small fraction of its prior size: from 60 contracts traded per week to one. However, due to regulatory pressure imposed by Basel III on the banking sector, safe haven CDS trading has been reported to grow at the same time (Klinger and Lando, 2015).

In light of the above market developments, liquidity of CDS is possibly driven by different factors in the two regimes. In the first half of the sample, especially during the financial crisis, liquidity risk could be an issue, as the European CDS market was not subject to clearing and thus these bilateral derivative transactions could have carried substantial counterparty credit risk. This could potentially mean that when market liquidity freezes, CDS liquidity might worsen in reaction. As opposed to this, in the second half of the sample, the level of liquidity and the depth of the sovereign CDs segment decreases substantially for riskier sovereigns. Given the complex market dynamics of the European sovereign CDS segment and the different regulatory changes affecting thereof, we leave empirically addressing this issue for future research.

1.5 Conclusion and extensions

This paper presents unique empirical evidence of selective default risk premium in inflation-linked sovereign bond (ILB) yields of Germany, France and Italy. Selective default is an event in which a sovereign issuer chooses not to meet obligations on a class of bonds, while servicing her other debt. We identify this effect from the difference of breakeven rates from country pairs. Differencing controls for common components, such as the effect of inflation expectations, monetary policy or interest rate risk. We find that the remaining part in breakeven rates is explained by two systematic risk factors, liquidity and sovereign credit risks - both within and across countries. We link these findings to the ILB-nominal puzzle, which shows that ILBs are underpriced relative to nominal bonds of the same issuer. We show that this underpricing is in part due to relative risk premia differences between nominal and inflation-linked debt: ILBs are less liquid, moreover investors perceive them to have higher credit risk during the financial and euro crises. This implies an implicit seniority and a subsequent convenience yield in nominal bonds.

Nonetheless, the method presented above has its limitations. First, the description of the results already pointed out specific features of the sample that make statistical inference more challenging. Our strategy to overcome this burden is to improve the estimation of the betas, either in a statistical or in an economic sense. The first could be performed by using a statistical method that allows me to benefit from the use of data at different frequencies as in Ghysels et al. (2007). By applying mixed data sampling or the MIDAS method, we could estimate and profit from having daily observations on market returns and credit risk, whereas we could simultaneously use these estimates with betas from the monthly illiquidity regressions. An alternative improvement is to impose structure on the beta estimation in an economic sense: make betas dependent on the asset characteristics, such as maturity or to construct another, higher frequency liquidity measure. Apart from the estimation, a natural extension would be to treat the identification as a trading strategy and to further explore the (limits to) its tradability and the risk-return characteristics.

Furthermore, and additional factor could be added to the analysis: the second PCA of liquidity measures suggests to capture flight-to-liquidity or safety phenomena. Adding this extra factor could help to further separate the liquidity and credit effects. Also, going in a similar direction, taking care of changing CDS liquidity and finding a viable proxy for expected credit risk could further improve the analysis. If such a proxy exists, another robustness test arose: cross-sectional SBEI regressions could be performed based on yields that are corrected for issuer credit risk, similarly to the corporate bond literature.

And finally, a related question to be answered is whether liquidity (and credit) proxies are affected and how the relative pricing of bonds changes in reaction to quantitative easing of the ECB. One could even go one step further to see how the spread of breakeven rates changes due to such unconventional monetary actions, and whether these markets price such events symmetrically in the nominal and inflation-indexed market segments. Studies to date, like Krishnamurthy and Vissing-Jorgensen (2011); Krishnamurthy et al. (2015), focus on the effects on different market segments separately, whereas one could look at the relative, potential lead-lag effects across indexed and nominal bonds.

Table 1.1
Descriptive statistics

The table presents descriptive statistics for variables used in the two-stage estimation. Panel A to C present variables for the analysis of German, French and Italian sovereign bond segments, respectively. All yields are quoted in annualized percentages terms, whereas yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields. Age and time-to-maturity are defined relative to the issue and maturity dates and are measured in days; while issued amount captures the size of a given issue in million EUR. Proportion of zero returns is the percentage of days with zero returns over a month. ILLIQ is the monthly ratio of absolute bond market returns over monthly aggregate trading volume, rescaled by 1 million EUR. Yields, volatilities and the zero returns measures are in percentages, age and time-to-maturity are measured in days. The data correspond to the sample period between July 2004 and February 2014.

Panel A: Descriptive statistics of German sovereign bonds

ILBs			**Nominal bonds**		
	Mean	St. Dev.		Mean	St. Dev.
ILB yield	0.479	0.972	Nominal yield	1.606	1.158
Yield volatility	0.000	0.199	Yield volatility	0.000	0.037
Age	992.5	733.6	Age	2317.0	2351.8
Time-to-maturity	2183.0	1033.1	Time-to-maturity	4219.0	3204.2
Issued amount (million)	13,600	1,498	Issued amount (million)	17,240	4,797
Proportion of zeros	0.407%	1.588%	Proportion of zeros	1.856%	8.331%
ILLIQ	1.176	1.649	ILLIQ	0.033	0.028

Panel B: Descriptive statistics of French sovereign bonds

ILBs			**Nominal bonds**		
	Mean	St. Dev.		Mean	St. Dev.
ILB yield	1.289	0.946	Nominal yield	2.926	1.272
Yield volatility	0.000	0.123	Yield volatility	0.000	0.039
Age	1694.9	1100.0	Age	2027.3	1788.5
Time-to-maturity	5032.8	3370.9	Time-to-maturity	4483.0	4002.6
Issued amount (million)	11,830	4,532	Issued amount (million)	24,770	7,852
Proportion of zeros	0.240%	1.362%	Proportion of zeros	0.451%	2.037%
ILLIQ	-	-	ILLIQ	-	-

Panel C: Descriptive statistics of Italian ILBs

ILBs			**2+Nominal bonds**		
	Mean	St. Dev.		Mean	St. Dev.
ILB yield	3.330	3.124	Nominal yield	4.157	1.270
Yield volatility	0.000	231.155	Yield volatility	0.007	1.576
Age	1137.7	913.1	Age	1899.5	1620.5
Time-to-maturity	3798.0	3100.7	Time-to-maturity	4200.4	2976.3
Issued amount (million)	12,420	3,954	Issued amount (million)	18,900	6,878
Proportion of zeros	0.194%	1.160%	Proportion of zeros	0.576%	4.360%
ILLIQ	1.049	1.475	ILLIQ	0.073	0.073

Table 1.2
Beta estimates

The table presents descriptive statistics for betas estimated from the time-series regression of bond returns on market, illiquidity and credit factors. The table consists of four major segments corresponding to the three countries and the aggregated portfolio betas, henceforth net betas. Panel A contains beta estimates from the ILB and nominal sectors of the German sovereign market, whereas Panel B and C do so for France and Italy, respectively. The forth part, Panel D, presents net betas that serve as a starting point for the breakeven based estimation. We estimated market, illiquidity and credit betas for all available German, French and Italian nominal and inflation-linked bond issues in the sample that spans the period between July 2004 and February 2014.

Panel A: German beta estimates

ILBs	Mean	St. Dev.	**Nominal bonds**	Mean	St. Dev.
Integrated market β	-0.2307	0.1939	Integrated market β	-0.2812	0.4833
Integrated illiquidity β	0.0002	0.0007	Integrated illiquidity β	-0.0008	0.0020
Integrated credit β	-0.0009	0.0011	Integrated credit β	0.0002	0.0017
Segmented market β	0.2669	0.3134	Segmented market β	0.2190	0.7175
Segmented illiquidity β	-0.0005	0.0003	Segmented illiquidity β	-0.0011	0.0012
Segmented credit β	-0.0011	0.0007	Segmented credit β	-0.0007	0.0010

Panel B: French beta estimates

ILBs	Mean	St. Dev.	**Nominal bonds**	Mean	St. Dev.
Integrated market β	0.1716	0.3144	Integrated market β	-0.0007	0.4549
Integrated illiquidity β	0.0021	0.0050	Integrated illiquidity β	0.0013	0.0023
Integrated credit β	0.0007	0.0039	Integrated credit β	0.0033	0.0040
Segmented market β	0.3874	0.3303	Segmented market β	0.1729	0.5330
Segmented illiquidity β	0.0038	0.0031	Segmented illiquidity β	0.0025	0.0025
Segmented credit β	0.0004	0.0021	Segmented credit β	0.0029	0.0025

Panel C: Italian beta estimates

ILBs	Mean	St. Dev.	**Nominal bonds**	Mean	St. Dev.
Integrated market β	0.1696	0.3648	Integrated market β	0.0788	0.3468
Integrated illiquidity β	-0.0048	0.0045	Integrated illiquidity β	-0.0020	0.0048
Integrated credit β	-0.0102	0.0070	Integrated credit β	-0.0055	0.0036
Segmented market β	0.3711	0.4327	Segmented market β	0.0742	0.1231
Segmented illiquidity β	0.0000	0.0027	Segmented illiquidity β	-0.0013	0.0033
Segmented credit β	-0.0073	0.0080	Segmented credit β	-0.0060	0.0039

Panel D: Net beta estimates

Net betas	Mean	St. Dev.
Net market β	-0.2100	0.4505
Net illiquidity β	-0.0005	0.0060
Net credit β	-0.0033	0.0057

Table 1.3
SBEI regressions: controlling for liquidity and sovereign risk

The table reports results from Equation 1.7, where liquidity and credit effect are directly identified from the difference of cross-country breakeven rates. The dependent variable is the spread on breakeven rates across two countries, whereas the net betas are constructed as a portfolio of individual bond betas from the market-level Fama-MacBeth regressions. Size of issue is measured as the natural logarithm of the amount issued, while age and time-to-maturity are defined relative to the issue and maturity dates, respectively. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions, where errors are adjusted for averaging and include a 12-lag Newey-West correction. The sample period is July 2004 until February 2014. Absolute values of the t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	(1)	(2)	(3)	(4)	(5)	(6)
Net illiquidity beta	-74.6564	8.0222	36.2303	5.1340	-0.2715	-2.6492
	(2.04)**	(0.36)	(1.52)	(0.31)	(0.01)	(0.1)
Net credit beta		-126.0375	-142.9042	-115.0918	-109.5895	-130.7344
		(5.40)***	(5.02)***	(3.89)***	(7.93)***	(4.52)***
Size of issue			-0.082766			
			(1.4)			
Age of issue				0.0137		
				(2.82)***		
Time-to-maturity					0.0039	
					(0.56)	
Net market beta						-0.5654
						(1.2)
Constant	-0.2922	-0.5114	-0.5753	-0.5759	-0.4790	-0.5339
	(5.72)***	(5.40)***	(8.33)***	(7.10)***	(4.21)***	(4.59)***
R2	0.31	0.57	0.79	0.79	0.64	0.7
N	866	866	866	842	842	866

Table 1.4
Risk-adjusted SBEI series

This table presents descriptive statistics the risk-adjusted spreads on breakeven series presented in Figure 1.5. We apply two adjustments, first only the effect of liquidity is considered, then following Equation 1.2, liquidity and credit risk premiums are both taken out. All figures are percentage yields. Panel A reports the average series across pairs for a given country pairing; whereas Panel B presents liquidity risk adjusted breakeven rates. Panel C also takes out the effect of credit risk. The data correspond to 6 pairs formulated between Germany and Italy, 5 pairs between Germany and France and 9 pairs from Italy and France for the sample period between July 2004 and February 2014.

Panel A: Average breakeven spreads

	Mean	St. Dev.	Min	Max
Germany vs. France	-0.360	0.287	-1.027	0.259
Germany vs. Italy	0.233	1.077	-0.666	4.009
Italy vs. France	-0.290	0.533	-2.727	0.240

Panel B: Liquidity risk-adjusted breakeven spreads

	Mean	St. Dev.	Min	Max
Germany vs. France	-0.147	0.288	-0.806	0.478
Germany vs. Italy	0.181	1.081	-0.761	3.956
Italy vs. France	-0.028	0.521	-2.417	0.545

Panel C: Liquidity and credit risk adjusted breakeven spreads

	Mean	St. Dev.	Min	Max
Germany vs. France	-1.873	0.264	-2.367	-1.361
Germany vs. Italy	-6.453	1.139	-7.774	-3.475
Italy vs. France	2.788	4.594	-4.993	7.569

Table 1.5

SBEI regressions: Trading volume-weighted asset level liquidity measures

The table reports results from the spread-on-basis regressions, where liquidity and credit effect are identified directly from the difference between matched basis series. The dependent variable is the maturity-matched breakeven yield difference across two countries, whereas net betas are constructed following Equation 1.10. Volume-weighting is based on the trading volume share of each bond market segment from the overall trading volume by primary dealers across the three countries. Size of issue is measured by the amount issued, while age and time-to-maturity are defined relative to the issue and maturity dates, respectively. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions where errors take into account the averaging and include a 12-lag Newey-West correction. The sample period is January 2006 until December 2013. Absolute values of the t-statistics are given in and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	(1)	(2)	(3)	(4)	(5)
Net illiquidity beta	7.5613	34.6524	20.6808	-152.6454	-3.6568
	(0.33)	(1.26)	(0.44)	(2.72)***	(0.14)
Net credit beta	-127.8355	-146.3623	-134.7653	3.194	-133.2498
	(5.44)***	(6.69)***	(3.47)***	-0.08	(4.60)***
Vol.-weighted size		0.0000			
		(6.45)***			
Vol.-weighted age			0.0048		
			(0.17)		
Vol.-weighted time-to-mat.				-0.2164	
				(3.74)***	
Net market beta					-0.5954
					(1.26)
Constant	-0.5147	-0.5407	-0.4077	-0.3597	-0.5447
	(5.32)***	(5.42)***	(2.83)***	(1.98)*	(4.70)***
R^2	0.58	0.80	0.81	0.74	0.70
N	818	818	685	685	818

Table 1.6
Liquidity and credit risks in pooled and integrated markets

This table reports pooled estimates from the second step of the Fama-MacBeth regressions with Eurozone-level risk factors. Therefore, the market factor is an equally-weighted average of all bonds across the three countries. Unlike in previous tables, this table presents regressions with bonds pooled together to estimate the effects of risk exposures. The dependent variables are the respective bond yields, and betas are estimated in the first step of the procedure as the loading on the common market, illiquidity risk and credit risk factors. We use the natural logarithm of the original issued amounts to capture the size of an issue. The average effect reported is the product of the estimated risk premium and the mean exposure to the specific risk factor. The calculated economic impact is defined as the interquartile spread coefficient times the difference between the betas that correspond to the first and third quartile in the cross-section of betas. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions, where errors are adjusted for averaging and include a 12-lag Newey-West correction. The sample period is July 2004 until February 2014. Absolute values of the t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Integrated mkt. beta	2.2425	2.2411	1.9554	1.9661	1.9107	2.1056	1.9194	2.0531
	(9.19)***	(9.06)***	(7.49)***	(7.48)***	(6.41)***	(7.61)***	(6.49)***	(6.97)***
Eurozone illiq. beta		105.8911	214.6902	220.2507	198.7834	206.9699	205.5075	218.9859
		(2.01)**	(4.23)***	(4.14)***	(4.10)***	(4.14)***	(3.90)***	(3.83)***
Eurozone credit beta			-106.0974	-106.7479	-122.7884	-125.1224	-124.764	-127.2105
			(4.47)***	(4.43)***	(6.27)***	(6.91)***	(6.29)***	(6.90)***
Size of issue				-0.0359			0.0124	0.0008
				(0.89)			(0.23)	(0.01)
Age					0.0197		0.0188	0.0179
					(3.97)***		(2.98)***	(2.59)**
Time-to-maturity						-0.0069		-0.0095
						(1.63)		(2.28)**
Constant	3.0914	3.0224	2.8477	3.6843	2.7604	2.9486	2.4509	2.8414
	(18.86)***	(18.61)***	(17.34)***	(3.57)***	(15.37)***	(15.54)***	(1.76)*	(1.97)*
R^2	0.29	0.43	0.54	0.55	0.56	0.57	0.57	0.58
N	10,894	10,894	10,894	10,894	9,600	9,620	9,600	9,600
Avg. effect of illiq.	-	-0.0805	-0.1631	-0.1674	-0.1510	-0.1573	-0.1562	-0.1664
Impact of illiq. risk	-	0.2752	0.5580	0.5724	0.5166	0.5379	0.5341	0.5691
Avg. effect of cr. risk	-	-	0.1735	0.1746	0.2008	0.2046	0.2040	0.2080
Impact of credit risk	-	-	-0.6738	-0.6780	-0.7798	-0.7947	-0.7924	-0.8079
Avg. effect of mkt. risk	-0.0485	-0.0485	-0.0423	-0.0425	-0.0413	-0.0455	-0.0415	-0.0444
Impact of market risk	1.2795	1.2787	1.1157	1.1218	1.0902	1.2014	1.0951	1.1714

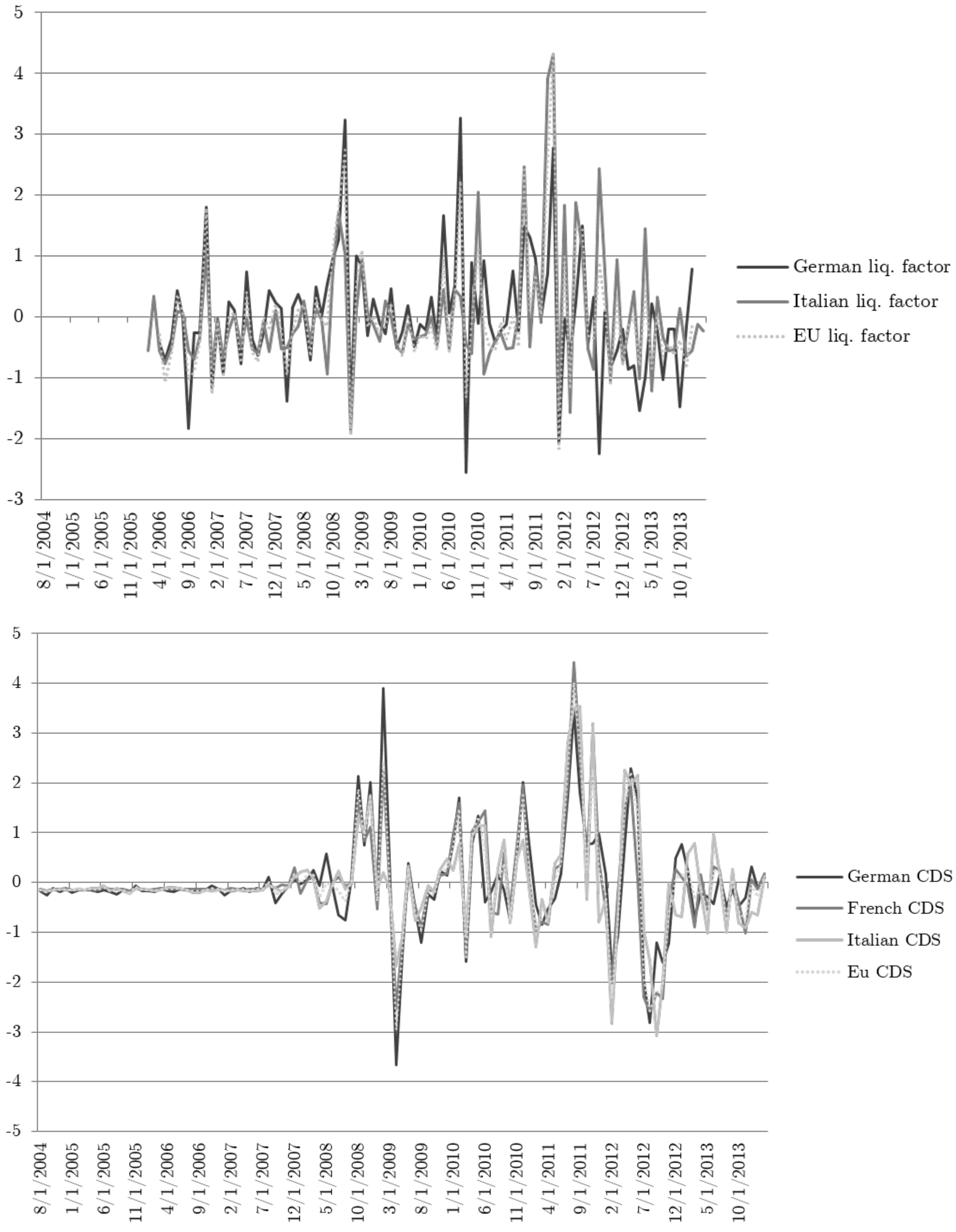


Figure 1.1 Illiquidity and credit risk factors

The figure depicts the The country and the integrated Eurozone risk factors: the upper panel shows the illiquidity factors, while the credit factors can be found in the lower panel. All factors are calculated as the standardized residuals from autoregressive processes imposed on the underlying liquidity and credit risk measures.

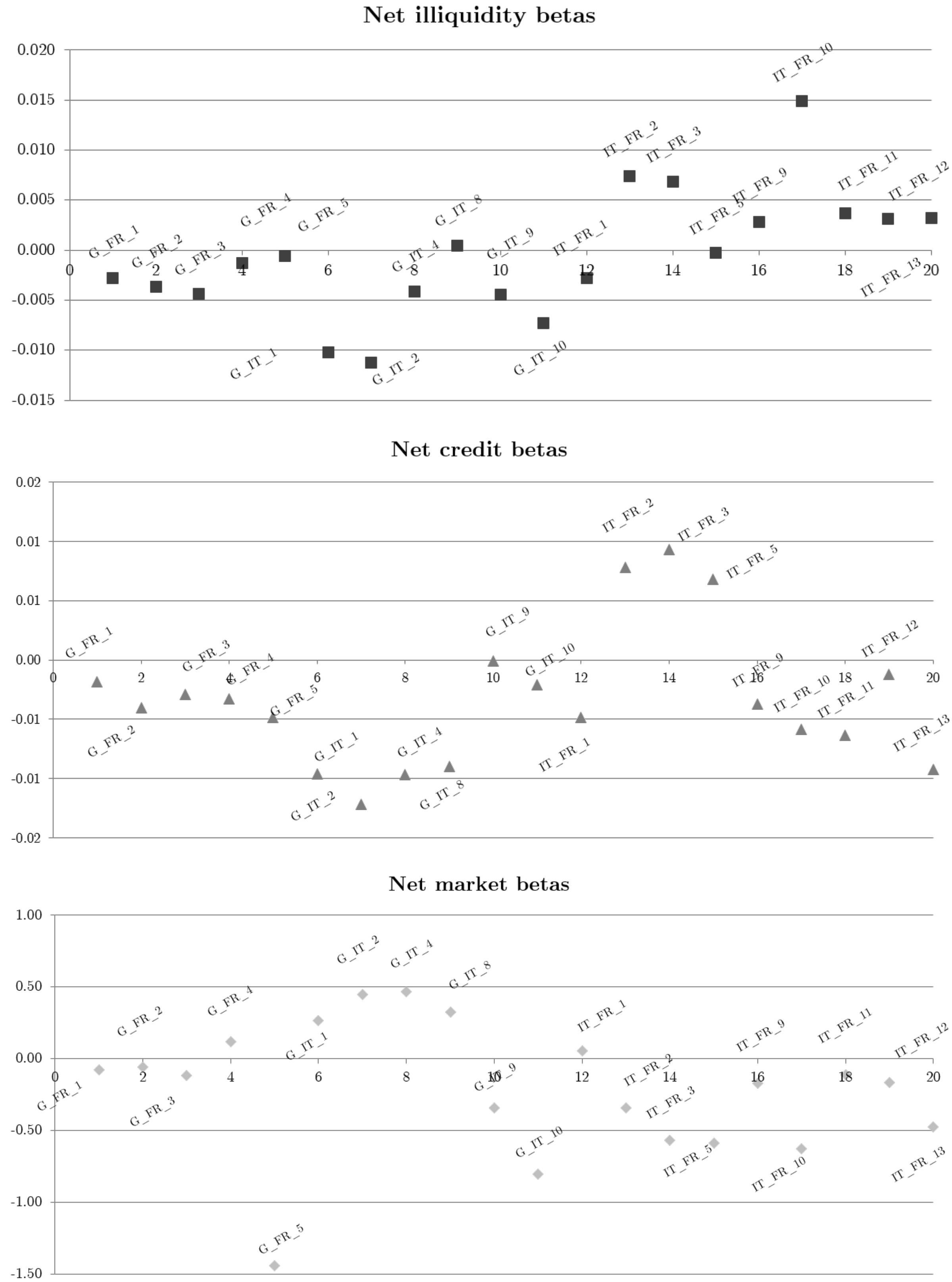


Figure 1.2 Net betas

The figure depicts the net betas following Equation 1.10. Net betas are essentially a portfolio of integrated betas of the four bonds in the spread-on-basis series. The upper panel shows the net illiquidity, the middle one the net credit, whereas the lower panel one the net market betas of the 20 spreads available in the sample. Labels indicate the pair of bases to which the individual betas correspond.

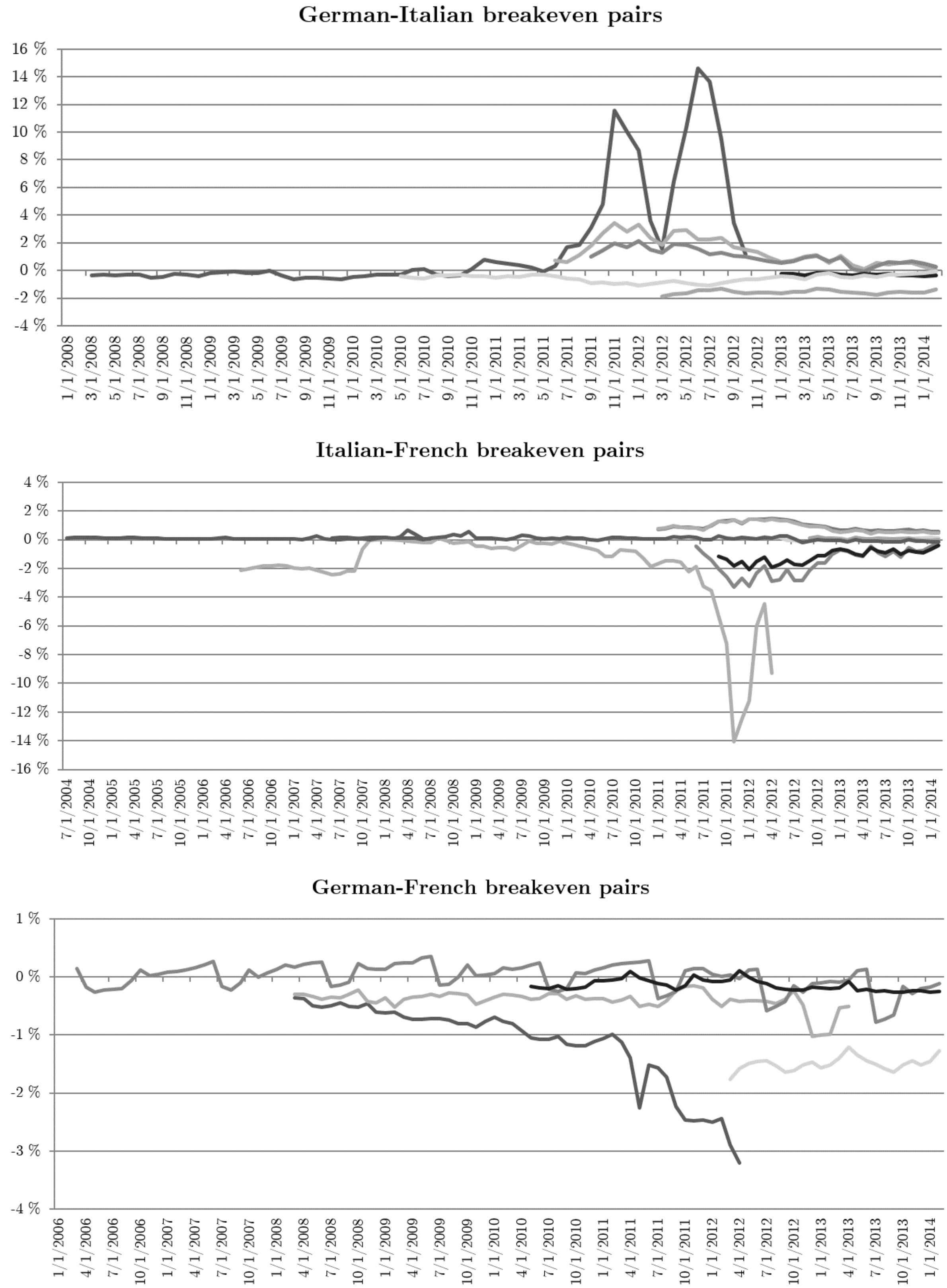


Figure 1.3 SBEI series

The figure depicts the percentage difference between maturity-matched breakeven rate pairs from different countries. In total there are twenty of such pairs: 6 pairs formulated between Germany and Italy (upper panel), 5 pairs between Germany and France (middle panel) and 9 pairs from Italy and France (lower panel).

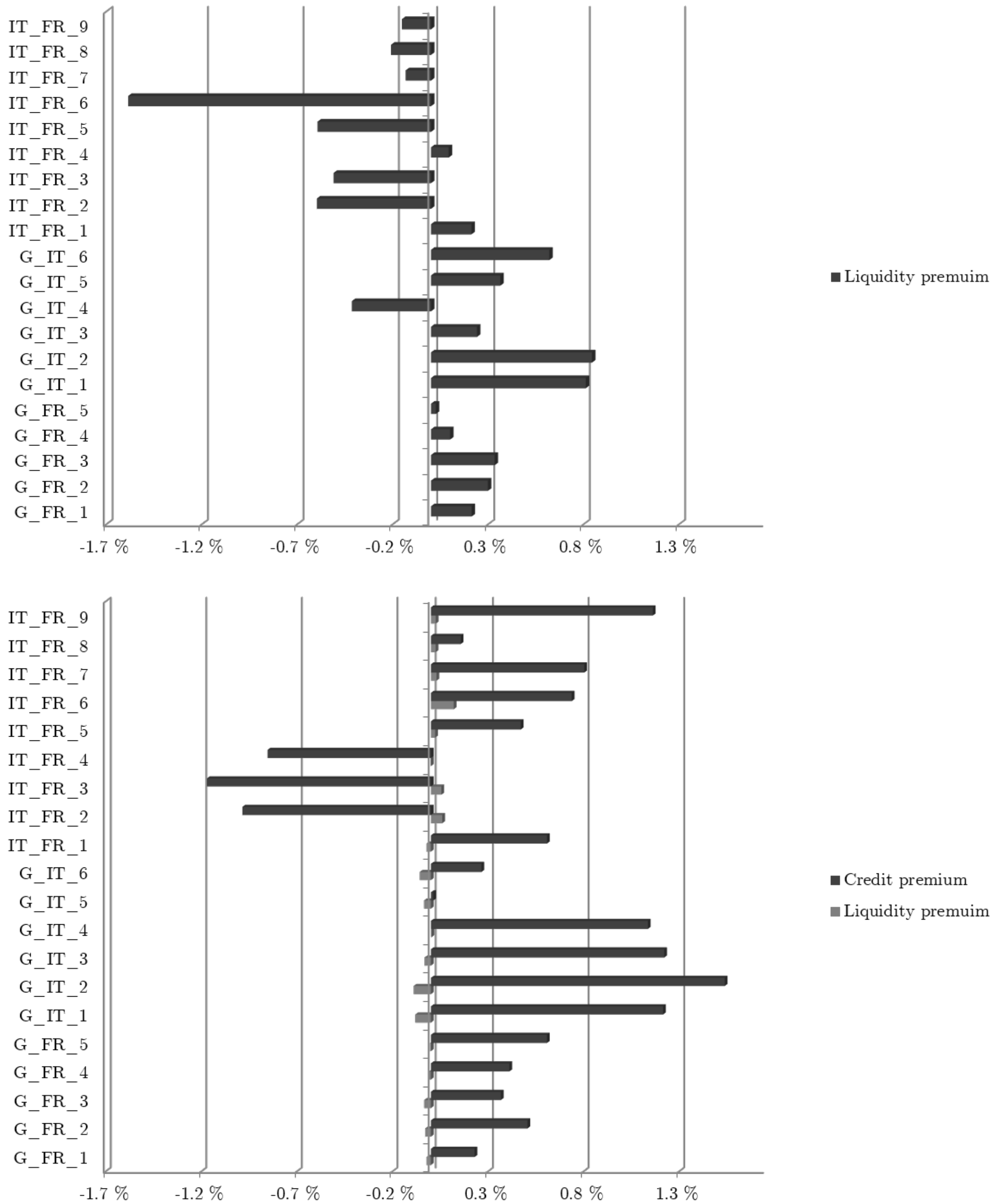


Figure 1.4 Risk premia identified in SBEI series

The figure depicts the size of percentage yield risk premium identified from spreads of cross-country breakeven pairs. The respective risk premium is the product of a certain pairs net beta and the market price of risk estimated from the cross-section of strategies. The upper panel shows estimates based on Column 1, whereas the lower panel considers estimates from Column 2 of Table 1.3.

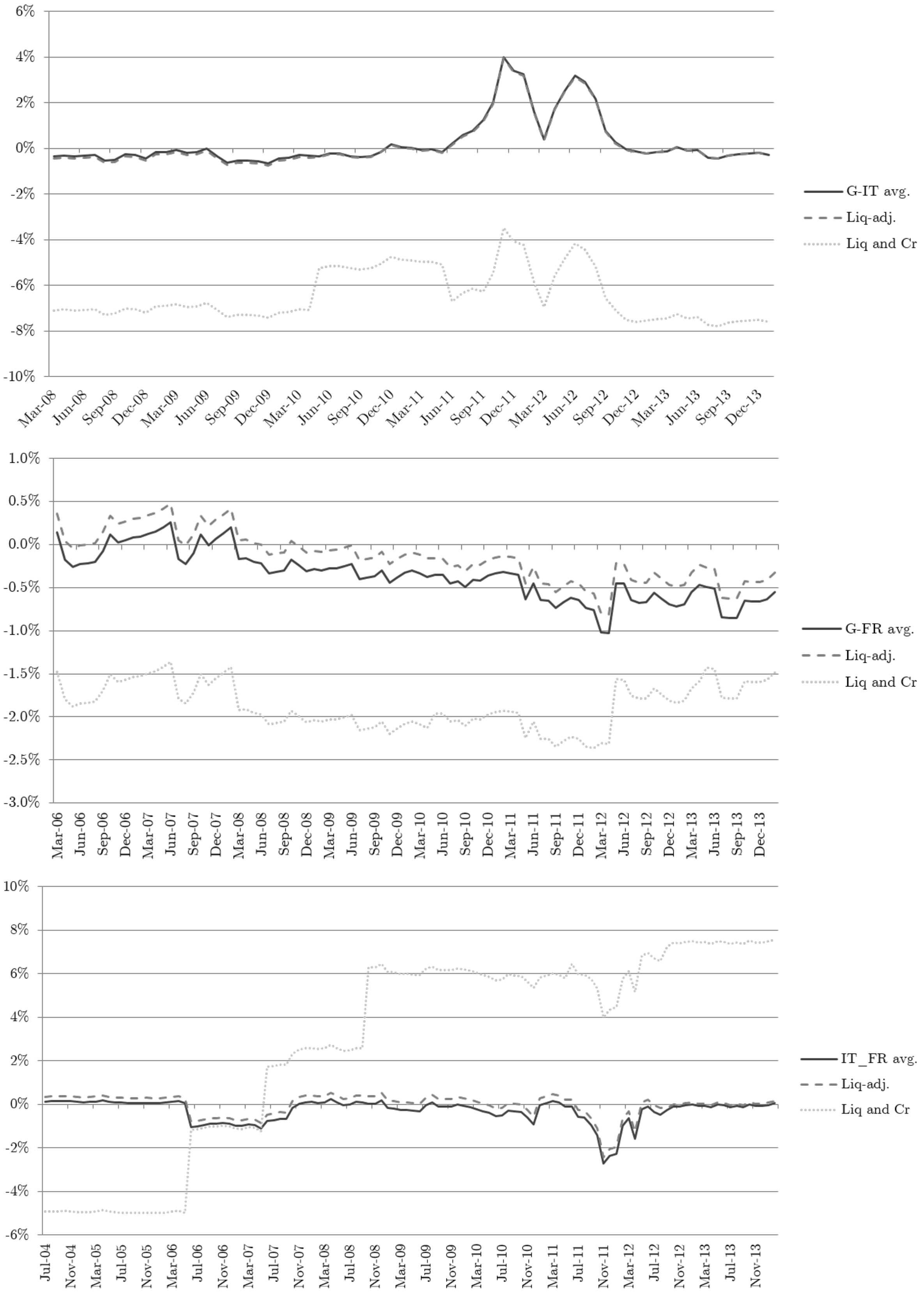


Figure 1.5 SRisk-adjusted SBEI series

The figure depicts the risk-adjusted cross-country breakeven spreads net of liquidity, or the sum of liquidity and credit effects. Solid lines denote the unadjusted, dashed the liquidity-adjusted, and dotted lines indicate series that are adjusted by the sum of liquidity and credit effects. The upper panel shows pairs from Germany and Italy, whereas the middle and lower panels depict German-French and French-Italian pairs, respectively.

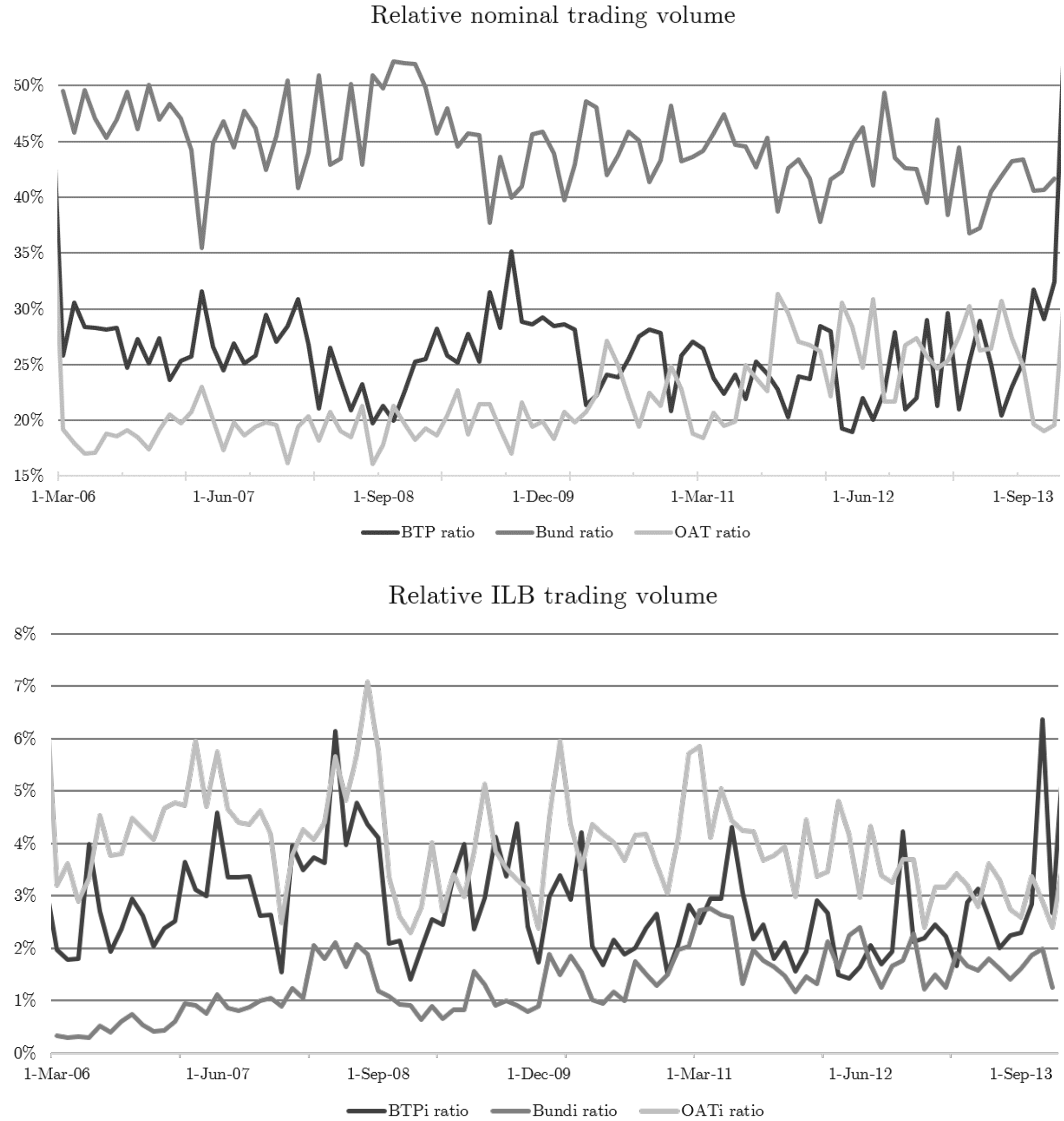


Figure 1.6 Relative trading volumes of each bond segment

The figure depicts the relative trading volume of each bond segment in the sample. The upper panel shows that of nominal bonds, whereas the lower panel contains the relative trading volume for the German, French and Italian HICP-linked bond segments. Note that the percentages of the two panels together add up to hundred percent.

1.A Appendix

This appendix shows that the direct identification of differential liquidity and selective default risk can not only be derived by using breakeven rates, but also by means of the swapped ILB-nominal basis based on Fleckenstein et al. (2014) and Fleckenstein (2013) . They define the basis or mispricing as the price difference between a nominal sovereign bond and a synthetic bond, which replicates the nominal cash flows. The latter is essentially an inflation swapped-indexed bond, whose cash flows are converted to fix payments exactly matching those of the corresponding nominal bond. Moreover, the maturities of the two nominal bonds are also matched.

To replicate this strategy, an investor buys an ILB issue and shorts a nominal bond at the same time. Additionally, she executes a zero-coupon inflation swap contract with the same maturity and notional amount as the ILB coupon and repeats this for each coupon and for the principal amount, which results in the execution of an entire swap portfolio. The rationale for swapping the bond is that the sum of the two cash flows is constant if they are linked to the same index and equal to the nominal coupon or principal. The investor also takes a small position in nominal principal STRIPS if there is disparity between the nominal swapped ILB cash flows. Based on this logic, she applies these steps to all coupon payments, which result in the successful conversion of the ILBs variable cash flow stream to the fixed one of the corresponding nominal bond.[19]

In sum, the investor short sells the nominal bond, buys the inflation-linked bond issue and holds portfolios of zero-coupon inflation swap contracts and nominal principal STRIPS. Absent liquidity and credit effects, these three components exactly replicating the fixed periodic coupons and the principal of the nominal bond should have the same price as the nominal bond. Finally, we calculate and compare the price of the replicating portfolio to that of the nominal bond. If in a frictionless world the resulting prices of these to securities differ, an arbitrage opportunity would arise. However, there is empirical evidence that both liquidity and credit risks affect European sovereign yield (Fontana and Scheicher, 2010; Palladini and Portes, 2011; Pelizzon et al., minga; Darbha and Dufour, 2014; Monfort and Renne, 2014). Therefore, this price discrepancy, henceforth called the swapped ILB-nominal basis, captured by this strategy, is most likely to be explained by the differences in liquidity and credit risk premia in the constituent asset prices.

[19] I only focus on bonds and swaps, although there is some empirical evidence that STRIPS are exposed to liquidity issues, see for instance Daves and Ehrhardt (1993); Jordan et al. (2000). Nevertheless, Bühler and Vonhoff (2011) show that principal STRIPS, the ones used in the strategy, are less affected. Consequently, we are less concerned that small positions that are further reduced in the spread-on-basis strategy, taken in these assets would carry a sizeable liquidity premium that could distort our results. Moreover, the potentially negative premium, which is in line with previous findings, would only work against me by widening the basis.

I construct the basis series in the spirit of Fleckenstein et al. (2014).[20] The time series evolution of these series is depicted on Figure 1.A.1, where the upper panel shows the country average s, whereas the lower panel displays the overall average across all 27 pairs of ILBs and nominal bonds across the three countries. The figure shows that the yield difference between the nominal and synthetic nominal bonds varies substantially over time and across the three countries. Negative values of the basis suggest that the nominal bond has a lower yield, thus higher price, than its replicating portfolio. A notable difference between these series and the one presented in Fleckenstein et al. (2014) is that unlike in the US, European series do switch signs over time. This means that the return varies over time depending on market conditions. In Germany and France, the series have a distinctive and large drop at the Lehman crisis, whereas the Italian series is the most volatile, exhibiting large swings around the financial and euro crises. The series plummet in late-2011 and mid-2012, potentially in reaction to ECB intervention, as discussed by Krishnamurthy et al. (2015) and Pelizzon et al. (minga).

This strategy is an appealing way of looking at the relative pricing of the constituent bonds. Fleckenstein et al. (2014) show that the basis cannot be due to differences in the tax treatment of nominal and indexed Treasuries, trading costs, repo financing, collateral value and pledgeability, eligibility of stripping or differences in their ownership structure. However, one could argue that the illiquidity of inflation swaps and the deflation floor of indexed bonds are accountable for the above price discrepancy. Unfortunately, these alternative explanations are not possible to formally eliminate within this setting.[21]

An economically relevant side product of developing the spread on breakeven strategy is that we improve upon the strategy of Fleckenstein et al. (2014) . We do so by exploiting the benefits of our cross-country sample: inspired by the difference-in-differences approach, we scrutinize the swapped ILB-nominal basis across countries by taking the difference of two such series, each coming from one of the countries in our sample. These series are depicted on Figure 1.3. For instance, one such strategy we could be looking at the difference between German and Italian bond pairs:

$$b_t^G - b_t^{\mathrm{IT}} = \left(y_{\mathrm{nom}}^G - y_{\mathrm{rep}}^G\right) - \left(y_{\mathrm{nom}}^{\mathrm{IT}} - y_{\mathrm{rep}}^{\mathrm{IT}}\right). \tag{1.A.1}$$

[20] A detailed technical description of the mechanics of the nominal-ILB mispricing or basis can be found in Fleckenstein (2013), who as part of the G7 countries looks at Italian, French and German sovereign bond markets.

[21] We know very little about the inflation swap market, where anecdotal evidence by citetFlemi2012b suggests that illiquidity can be severe; and the deflation option cannot be hedged by inflation options either. The asset that is closest to an inflation option is the inflation spread option that takes the spread between two inflation indices and pays if the spread is positive. However, these assets are rarely traded, thus they are illiquid, and carry a sizeable counterparty risk premium (Kerkhof, 2005). Moreover, the value of the deflation option varies across bond maturities, as in Grishchenko and Huang (2013); Christensen and Gillan (2013).

The basis can be seen as the return of the Fleckenstein et al. (2014) strategy and thus defined as a portfolio consisting of a nominal issue, an inflation-indexed bond and inflation swaps. The return on such a portfolio is the sum of the returns on the constituent assets:

$$\begin{aligned}\mathbb{E}\left[R_t^G - R_t^{\mathrm{IT}}\right] &= \left(R_{\mathrm{nom}}^G - R_{\mathrm{rep}}^G\right) - \left(R_{\mathrm{nom}}^{\mathrm{IT}} - R_{\mathrm{rep}}^{\mathrm{IT}}\right) \\ &= \left[R_{\mathrm{nom}}^G - \left(R_{\mathrm{ILB}}^G + R_{\mathrm{swap}}\right)\right] - \left[R_{\mathrm{nom}}^{\mathrm{IT}} - \left(R_{\mathrm{ILB}}^{\mathrm{IT}} + R_{\mathrm{swap}}\right)\right]. \qquad (1.A.2)\end{aligned}$$

denotes the return on a German nominal bond, whereas refers to that of the cash flow replicating portfolio that consists of a maturity matched ILB issue and a portfolio of swap components. IT superscripts refer to the same assets from a similar Italian bond pair.

Although the above strategy focuses on the relative pricing of nominal and inflation-indexed bonds, due to the cash flow matching at all coupon dates, portfolios of inflation swaps and STRIPS are required for the exact cash flow replication. To construct the above strategy, we start with certain assumptions on asset positions within the strategy that will be later relaxed to get to a more general case. We presume first, that:

Nominal bonds have the same coupon as the swapped indexed coupon. This applies to the principal payments too. The swapped indexed coupons are equal across the two countries; thus the swap positions are virtually the same. This happens if the indexed bonds have the same coupon rate and coupon payment structure, for instance annual coupons. Nominal and indexed bonds have the same maturity date.

If all three of these conditions applied, the swap portfolios in the German and Italian bases would be the same and no STRIPS positions were required. Then we substitute expected returns by their forward-looking proxy, by yield-to-maturity, the following relationship arises:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] &= \left[R_{\mathrm{nom}}^{\mathrm{G}} - \left(R_{\mathrm{ILB}}^{\mathrm{G}} + R_{\mathrm{swap}}\right)\right] - \left[R_{\mathrm{nom}}^{\mathrm{IT}} - \left(R_{\mathrm{ILB}}^{\mathrm{IT}} + R_{\mathrm{swap}}\right)\right] \\ &= \left(R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{ILB}}^{\mathrm{G}}\right) - \left(R_{\mathrm{nom}}^{\mathrm{IT}} - R_{\mathrm{ILB}}^{\mathrm{IT}}\right), \qquad (1.A.3) \\ &= \mathbb{E}\left(R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{ILB}}^{\mathrm{G}}\right) - \mathbb{E}\left(R_{\mathrm{nom}}^{\mathrm{IT}} - R_{\mathrm{ILB}}^{IT}\right) \approx \left(y_{\mathrm{nom}}^{\mathrm{G}} - y_{\mathrm{ILB}}^{\mathrm{G}}\right) - \left(y_{\mathrm{nom}}^{\mathrm{IT}} - y_{\mathrm{ILB}}^{\mathrm{IT}}\right), \qquad (1.A.4)\end{aligned}$$

where is the return on a portfolio of different swap positions within the respective strategy. These positions depend on the coupon difference between the nominal and indexed bond as well as on the indexation or reference inflation of the ILB. Note that if the swap portfolios coincide, which holds by assumption, their respective returns cancel out. This leaves me with returns from the bond positions that, in expectation, can be proxied by the differences in their yields.

Equation 1.A.4 shows that by differencing two basis series we get the difference of two breakeven rates. In particular, this is an improvement of Fleckenstein et al. (2014), because differencing successfully eliminates the confounding effects in their strategy. It cancels out any common factors, which offers a clean way of testing the drivers of the relative pricing of indexed and nominal bonds in an international setting. First, the deflation option that applies to all bonds in the sample identically is fully hedged out: bonds in our sample are all indexed to the same inflation index, the HICP index. Furthermore, they also carry the same optionality regarding negative inflation their principal is protected against deflation but not the individual coupon payments. Consequently, after differencing the floors effect is fully diminished from this new strategy, so are any common market factors for the same reason. Secondly, because in all three markets the same HICP inflation swaps are traded, the inflation components also cancel out whenever the inflation indexed and nominal coupon rates are the same in both countries. If this does not hold, these positions are still negligibly small and the magnitude depends on the coupon difference of the constituent nominal and inflation bonds across the two countries.

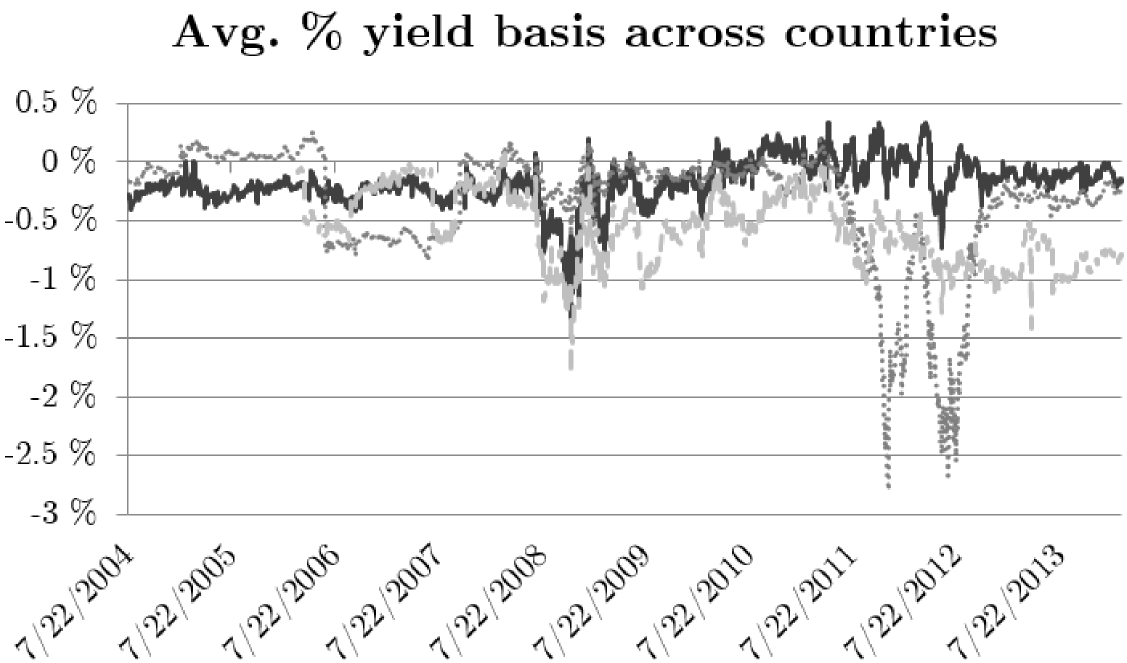
The two crucial assumptions for the inflation swap positions to cancel out are 1) either the coupons of the nominal or the synthetic bonds coincide in which case the swap positions cancel out within countries; or 2) we need the swapped ILB coupons to be equal across countries. Note that the first assumption makes sure that no positions in STRIPS are required. However, in reality maturities of bond pairs and index linked coupons are unlikely to match. If we relax the second and third assumptions, we get the following:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] &= \left(R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{rep}}^{\mathrm{G}}\right) - \left(R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{rep}}^{\mathrm{G}}\right) \\ &= \left[R_{\mathrm{nom}}^{\mathrm{G}} - \left(R_{\mathrm{ILB}}^{\mathrm{G}} + R_{\mathrm{swap}}^{\mathrm{G}}\right)\right] - \left[R_{\mathrm{nom}}^{\mathrm{IT}} - \left(R_{\mathrm{ILB}}^{\mathrm{IT}} + R_{\mathrm{swap}}^{\mathrm{IT}}\right)\right] \\ &= R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{nom}}^{\mathrm{IT}} + R_{\mathrm{ILB}}^{\mathrm{IT}} - R_{\mathrm{ILB}}^{\mathrm{G}} - \left(R_{\mathrm{swap}}^{\mathrm{G}} - R_{\mathrm{swap}}^{\mathrm{IT}}\right). \qquad (1.A.5)\end{aligned}$$

Where the superscript of the swap portfolio return indicates that the underlying swap positions differ depending on which countrys mispricing series they are coming from. This difference determines the residual swap position, which depends on the coupon difference between the ILBs, and is likely to be small. Taking the generalization one step further, another departure from the ideal case is when swapped ILB coupons do not match the nominal ones. By relaxing the first assumption we get the closest to reality, where we need to introduce nominal principal STRIPS positions into the portfolio:

$$\begin{aligned}\mathbb{E}\left[R_t^{\mathrm{G}} - R_t^{\mathrm{IT}}\right] &= \left(R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{rep}}^{\mathrm{G}}\right) - \left(R_{\mathrm{nom}}^{\mathrm{IT}} - R_{\mathrm{rep}}^{\mathrm{nom}}\right) \\ &= \left[R_{\mathrm{nom}}^{\mathrm{G}} - \left(R_{\mathrm{ILB}}^{\mathrm{G}} + R_{\mathrm{swap}}^{\mathrm{G}} + R_{\mathrm{STRIPS}}^{\mathrm{G}}\right)\right] - \left[R_{\mathrm{nom}}^{\mathrm{IT}} - \left(R_{\mathrm{ILB}}^{\mathrm{IT}} + R_{\mathrm{swap}}^{\mathrm{IT}} + R_{\mathrm{STRIPS}}^{\mathrm{IT}}\right)\right] \\ &= R_{\mathrm{nom}}^{\mathrm{G}} - R_{\mathrm{nom}}^{\mathrm{IT}} + R_{\mathrm{ILB}}^{\mathrm{IT}} - R_{\mathrm{ILB}}^{\mathrm{G}} - \left(R_{\mathrm{swap}}^{\mathrm{G}} - R_{\mathrm{swap}}^{\mathrm{IT}}\right) - \left(R_{\mathrm{STRIPS}}^{\mathrm{G}} - R_{\mathrm{STRIPS}}^{\mathrm{IT}}\right). \end{aligned} \tag{1.A.6}$$

The size of the STRIPS position depends on the difference between nominal and swapped index coupons. They are in general small when regular coupon payments occur, however, they might become sizeable for the principal payment. In order to circumvent large STRIPS positions, we looked for the closest possible match in terms of both maturity and tenor when maturity-matching nominal and indexed bonds in the sample. The benefit of matching tenors comes from the fact that bonds issued in similar economic environment (e.g.: low inflation) tend to have fairly similar coupons. This shrinks the position that one has to take as the difference of the nominal coupon and the swapped-ILB coupon.



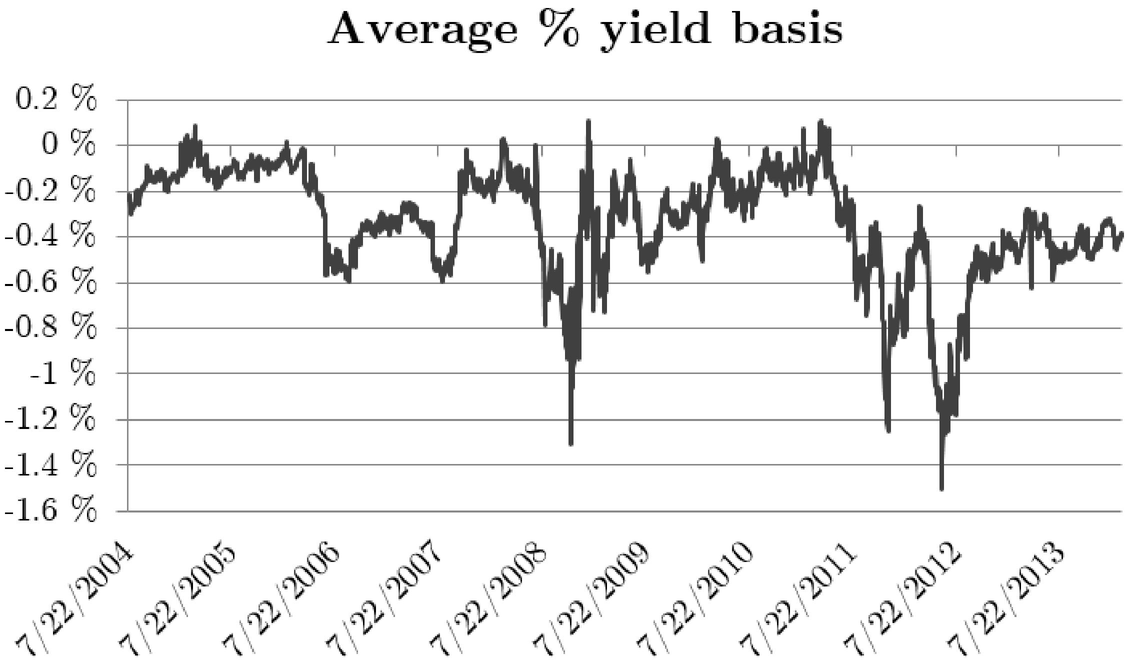


Figure 1.A.1 Aggregate basis series

The figure depicts the aggregate basis series defined as the yield difference of a nominal issue and its replicating portfolio following Fleckenstein et al. (2014). The upper panel shows the country-level aggregate series, where Germany is displayed in medium gray jagged, France in solid black and the Italian series is dotted and light gray. The aggregation takes places across all maturity-matched bond pairs of a given country. The lower panel shows the average of these three series.

1.B Appendix

This appendix contains graphs of the different liquidity and credit measures applied in the analysis. Most of these serve as a basis for constructing the principal components that are direct inputs for the risk factors. We also include the graph depicting the time-series dynamics of the sovereign CDS prices.

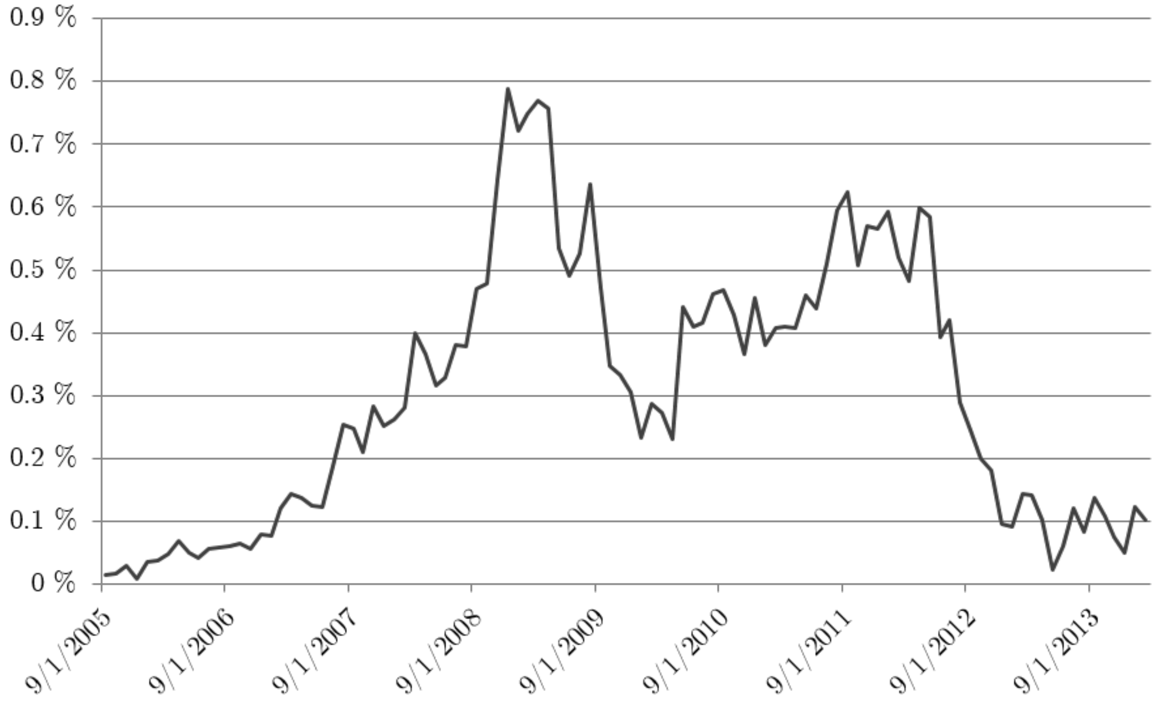


Figure 1.B.2 The KfW spread

The figure depicts the aggregate basis series defined as the yield difference between a nominal issue and its replicating portfolio following Fleckenstein et al. (2014). The upper panel shows the country-level aggregate series, where Germany is displayed in medium gray jagged, France in solid black and the Italian series is dotted and light gray. The aggregation takes places across all maturity-matched bond pairs of a given country. The lower panel shows the average of these three series.

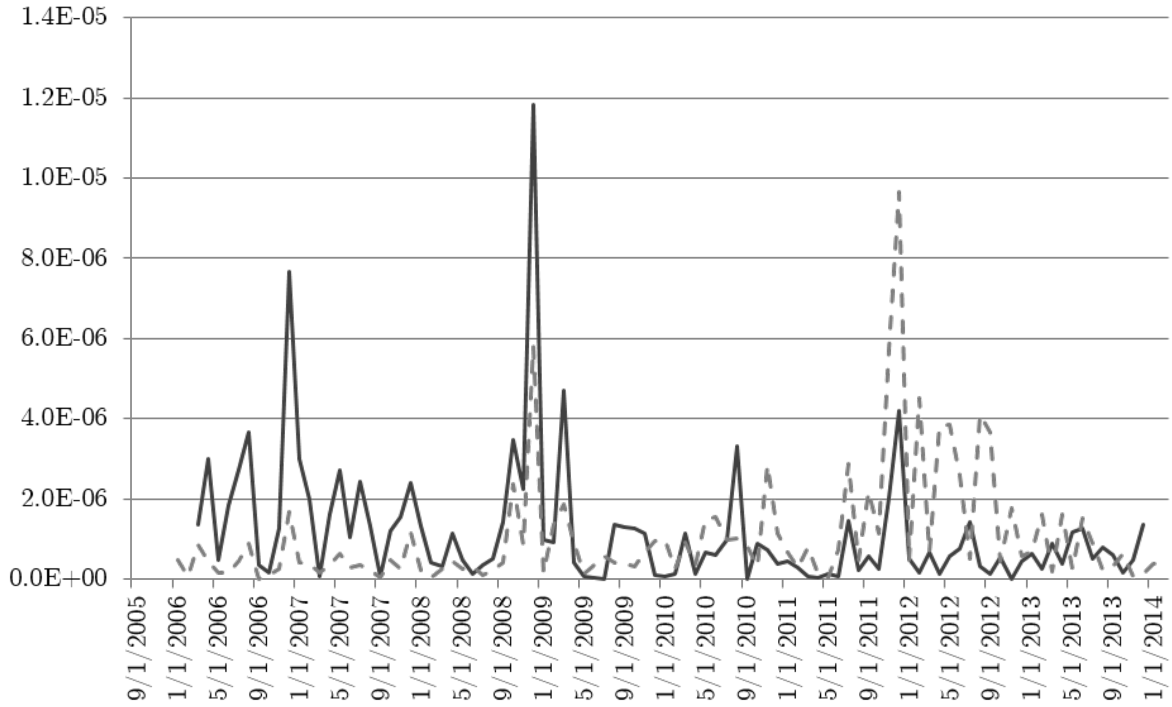


Figure 1.B.3 The ILLIQ measures for the ILB sectors

The figure depicts the time evolution of ILLIQ measures for German (solid line) and Italian (dashed line) ILBs

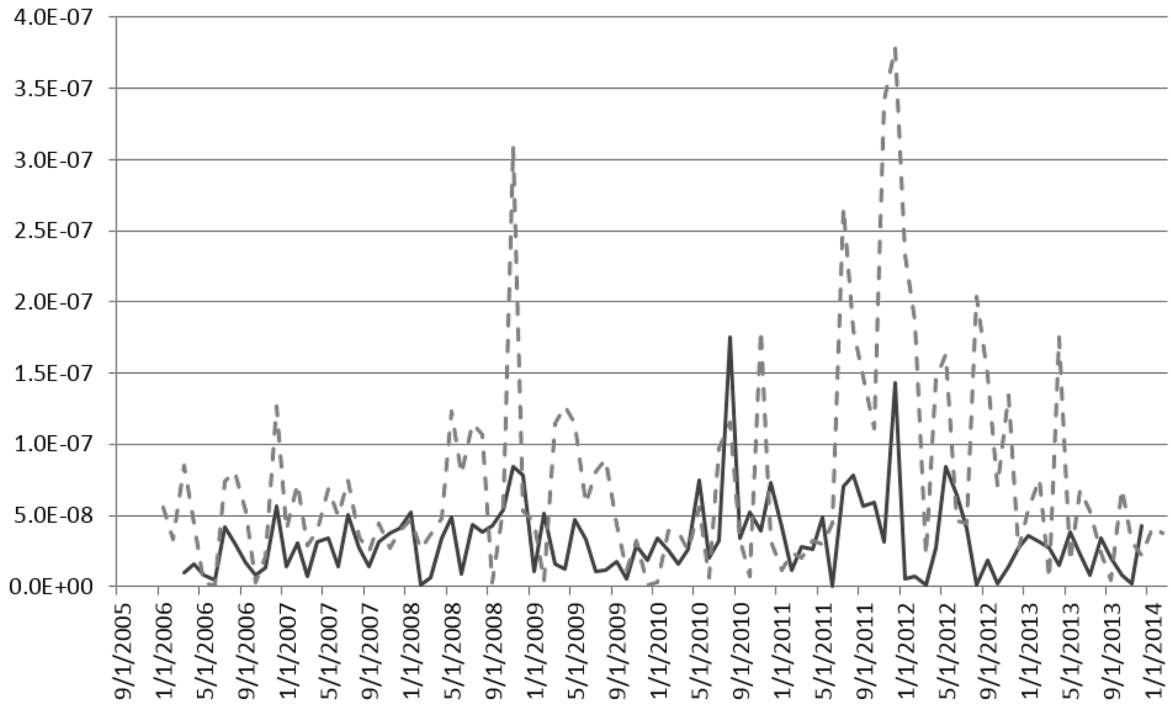


Figure 1.B.4 The ILLIQ measures for the nominal sectors

The figure depicts the time evolution of ILLIQ measures for German (solid line) and Italian (dashed line) nominal bonds.

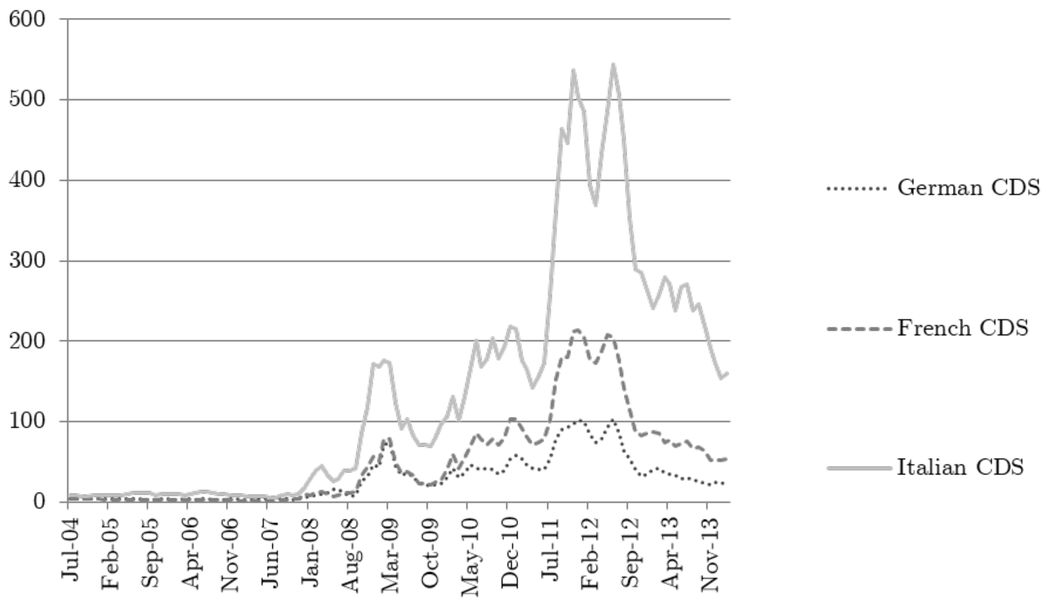


Figure 1.B.5 Sovereign CDS spreads

The figure depicts the time evolution of country-level CDS series. The solid line denotes Italy, whereas German and French series are dotted and dashed, respectively.

Chapter 2

The missing piece of the puzzle: Liquidity premiums in inflation-indexed markets

2.1 Introduction

A substantial literature has studied liquidity effects in government bond markets. Several studies show that, at least for the US, nominal bonds are very liquid and thus exhibit only small liquidity premiums.[1] Much less is known about liquidity effects in the markets of inflation-indexed products, such as inflation-indexed bonds (TIPS) and inflation swaps. In this paper we perform a detailed study of liquidity effects in these markets.

Understanding these liquidity effects is important for several reasons. First, liquidity effects directly matter for the relative pricing of nominal and indexed bonds as well as the breakeven inflation rate implied by these prices. Similarly, liquidity effects in inflation swaps affect the inflation expectations that can be extracted from these swap prices. Finally, recent work by Fleckenstein et al. (2014) finds evidence for large price differential of nominal bonds relative to a replicating strategy of indexed bonds and inflation swaps, and based on additional analyses they argue that this is due to mispricing of the indexed bonds. We assess whether part of this price differential is not mispricing but due to differences in liquidity premiums in the underlying asset markets.

This chapter is based on joint work with Joost Driessen and Theo Nijman. I would like to thank Dion Bongaerts, Frank de Jong, Alexander de Roode, Stefan Ruenzi and the seminar and conference participants of Tilburg Univeristy, Netspar Conference: Illiquid Investments and Robust Portfolio Choice, Netspar International Pension Workshop, Netspar Pension Day.

[1] See for instance Krishnamurthy (2002), and Longstaff (2004), among many others.

Our paper has two main contributions. The first contribution is that we show that in both index-linked bond markets and inflation swap markets liquidity is an important determinant of prices. We do so by estimating a model with both a liquidity risk factor and asset-specific liquidity characteristics. The use of a liquidity risk factor is inspired by Pastor and Stambaugh (2003) and Acharya and Pedersen (2005). To estimate the effect of liquidity risk, we measure an asset's exposure to our non-traded liquidity factor. In addition to this liquidity risk exposure, the level of liquidity is proxied by asset-level characteristics, following Krishnamurthy (2002) and Houweling et al. (2005). We also study liquidity effects in nominal bonds in a similar way, so that in total we analyze liquidity premiums in three markets.

We conduct our analyses based on two alternative assumptions – we either propose the three markets being segmented, such that prices are independently determined, or integrated markets.[2] In our benchmark specification, corresponding to segmented markets, we find in the TIPS market the effect of illiquidity risk is dominated by asset characteristics such as age and the size of an issue, together carrying a sizable premium of 32.68 basis points estimated at the monthly frequency. This effect means that higher age and lower size together increase the yield of a TIPS issue, which translates into a lower price. Age and size are bond characteristics that capture liquidity of an issue as in Houweling et al. (2005). They argue that the more time passes since issuance, the more likely an issue gets locked up in buy-and-hold investor portfolios which decreases liquidity. On the other hand, larger bond issues tend to be more liquid. As for inflation swaps, we find that illiquidity risk is priced yet the premium and the implied economic effect, a monthly 1.65 basis points, is small. Finally, we find a small liquidity risk premium in nominal bond markets.[3] These novel results are robust to the inclusion of various controls and to shifting to the proposition of integrated markets. Then results regarding TIPS and nominal Treasuries are akin to the benchmark case, while for swaps the price of illiquidity risk is negative and twice as large relative to the benchmark case, -3.41 basis points per month.

Our second main contribution is that we scrutinize whether the above diversity in exposure to liquidity and liquidity risk could explain the persistent difference in relative bond prices, as documented in Fleckenstein et al. (2014). They show that there exist substantial price differences between a nominal Treasury bond and its synthetic counterpart -

[2]Under the assumption of market segmentation, our estimates are bounded by the mathematics of factor models: the average of our market betas is one and that of the liquidity betas is zero by construction, whereas in the integrated case there are no such restrictions. Consequently, we define our benchmark results as that corresponding to the segmented case. This specification is more conservative and is also less prone to biased estimates due to omitted variables.

[3]The magnitude of our estimate is similar to the on-the-run spread of Krishnamurthy (2002), yet smaller than the premium estimated by Fontaine and Garcia (2012).

a swapped TIPS issue. We provide evidence that when controlling for liquidity, a large part of this price differential disappears.

By showing the importance of liquidity for the three markets we contribute to the long-standing literature on the effect of liquidity on asset prices (Amihud and Mendelson (1986), Amihud (2002), Bekaert et al. (2007), and Bongaerts et al. (2011) among many others). More specifically, we follow the footsteps of Pastor and Stambaugh (2003) and Acharya and Pedersen (2005) to show that liquidity risk is priced and provide novel evidence for Treasury bonds and inflation swaps. Moreover, we are also among the first ones to examine the effect of liquidity on inflation swaps, an important inflation derivative market (see e.g.: Chen et al. (2007), Bongaerts et al. (2011), and Tang and Yan (2007) on liquidity of other derivative markets, and Kerkhof (2005), and Fleming and Krishnan (2012) on inflation swaps specifically).

The paper also contributes to the literature on inflation-indexed bond pricing; more specifically we enrich the strand of papers that tackles liquidity of TIPS. For instance, D'Amico et al. (2010), Gürkaynak et al. (2010), and Haubrich et al. (2012) propose term structure models where they incorporate potential illiquidity of real bonds. Besides, Fleming (2003), Fleming and Sporn (2012), and Fleming and Krishnan (2012) focus on the microstructure characteristics of TIPS. Others like Campbell et al. (2009), Christensen and Gillan (2011), Pflueger and Viceira (2011, 2015), Fleckenstein et al. (2014) and Fleckenstein (2013) specifically focus on the relative pricing of nominal and indexed Treasuries, and the breakeven rate, which is the yield difference between these two securities. Our study deepens the understanding on this matter by examining the effect of both liquidity and liquidity risk of these securities. Our work is distinguished from prior literature by the empirical strategy that simultaneously examines the no-arbitrage relation between TIPS and nominal Treasures and the liquidity characteristics of constituent asset markets within the framework of a tradable strategy.

The approach of this paper is the closest related to Pflueger and Viceira (2015) and Fleckenstein et al. (2014). Similarly to Pflueger and Viceira, we identify risk premiums and liquidity effects in bond yields.[4] Yet unlike their paper, we do not incorporate a time varying behavior of risk premiums and the consequent return predictability in our analysis, as our primary objective is to quantify the effect of liquidity and liquidity risk on expected returns of TIPS, nominal Treasuries and inflation swaps. We also aim to answer to what extent the mispricing, found by Fleckenstein et al. (2014), is an artifact of liquidity premiums. We view our work as an extension of Fleckenstein et al. (2014) and Fleckenstein (2013), as we investigate the same no-arbitrage relationship between

[4]Like in Cochrane and Piazzesi (2002, 2005), Vayanos (2004), and Buraschi and Jitsov (2005) among others.

nominal and indexed Treasuries. Fleckenstein et al. (2014) provide evidence that part of this mispricing is caused by slow moving capital. (see Gromb and Vayanos (2002), Mitchell et al. (2007), Brunnermeier and Pedersen (2009), Duffie (2010) and Ashcraft et al. (2010)), while we show that market liquidity premiums are important to understand the price differences of indexed and nominal bonds.

In conclusion, our work differs from these abovementioned papers in four main aspects. First, unlike Pflueger and Viceira (2015) and Fleckenstein et al. (2014), we estimate the effect of liquidity based on a factor model that allows us to differentiate between liquidity premiums stemming from level and risk effects. Second, our primary liquidity proxy, the ILLIQ measure in Amihud (2002) does not rely on any implicit assumptions on the relative liquidity of nominal and indexed bonds, like in Pflueger and Viceira (2015) who assume nominal Treasuries to be perfectly liquid. In line with previous empirical evidence, we allow nominal Treasuries to also carry compensation for liquidity risk or a convenience yield Krishnamurthy (2002), Longstaff (2004), Krishnamurthy and Vissing-Jorgensen (2010)). Third, we study the liquidity effects in inflation swap markets. And finally, we assess to what extent the price differential between indexed and nominal can be explained by liquidity premiums.

Our data are an extended and updated version of Fleckenstein et al. (2014): we include a larger cross-section of bond issues and longer span (July 2004 - December 2011). The data consist of maturity-matched indexed and nominal issues and zero coupon inflation swaps. We complement this data with input for liquidity proxies and additional controls from multiple sources.[5]

To show that in both index-linked bond markets and inflation swap markets liquidity is an important determinant of prices, we estimate a model with both a liquidity risk factor and asset-specific liquidity characteristics. We define the risk factor as the surprise or unexpected illiquidity which is captured by the residual from an autoregressive process imposed on the illiquidity measure. This approach is similar to Acharya and Pedersen (2005), while we use the ILLIQ measure of Amihud (2002). To estimate the effect of liquidity risk, we measure an asset's exposure to our non-traded liquidity factor. We do this by following a two-stage Fama-MacBeth procedure in which we estimate market and liquidity betas from excess returns in the first stage. In the second stage we run repeated-cross sectional regressions of yields on these betas, the level of liquidity proxied by asset-level characteristics and additional controls. We estimate betas and risk loadings in each market separately and under the assumption of integrated or segmented markets. Given the liquidity measures at hand, we are able to measure the covariation of a security's

[5]Bloomberg, Datastream, Primary Dealer Historical Search database of the Federal Reserve of New York, Kenneth French's website and St. Louis Fed's FRED database.

return with the market liquidity, the same covariance as in Pastor and Stambaugh (2003).

We also study liquidity effects in nominal bonds in a similar way, which allows us to inspect the relative pricing of indexed and nominal Treasuries, following Fleckenstein et al. (2014). The idea behind their TIPS-Treasury arbitrage is that an investor matches the maturities and payoffs of a nominal bond issue and its synthetic counterpart. The latter is essentially an inflation swapped-indexed bond, whose cash flows are converted to fix payments exactly matching those of a corresponding nominal bond. We incorporate the effect of liquidity by adjusting the yields of nominal and indexed Treasuries by the estimated premiums of the benchmark cases. We also include the effect of liquidity on inflation swap positions concerning every coupon payment within the strategy.

The remainder of the chapter is organized as follows. Section 2.2 and 2.3 discuss the theory and the methodology of this study, whereas Section **??** describes the data and the constituent asset markets. In Section 2.4 we present our empirical results, and finally, Section 2.6 concludes.

2.2 Pricing of liquidity in the Treasury bond, TIPS and inflation swap markets

In this section we explain our empirical strategy to examine the effects of liquidity and liquidity risk on prices of nominal and indexed Treasuries alongside with inflation swaps. We base our empirical identification strategy on previous empirical findings: Amihud and Mendelson (1986) and Amihud (2002) show that average liquidity is priced on stock markets, both in the cross-section of stocks and over time. Pastor and Stambaugh (2003) find that return sensitivity to market liquidity is priced, whereas Acharya and Pedersen (2005) present a model that disentangles three sources of liquidity risk - each being priced in the market.

We aim to test the following relationship between excess returns and liquidity:

$$\mathbb{E}(R_{i,t} - R_{\mathrm{f},t}) = \mathbb{E}(Liq_{i,t}) + \lambda_{\mathrm{LIQ}}\beta_{\mathrm{LIQ},i} + \lambda_{\mathrm{MKT}}\beta_{\mathrm{MKT},i}, \tag{2.1}$$

where $\mathbb{E}(Liq_{i,t})$ is the unconditional expectation of the liquidity measure corresponding to asset i at time t, that aims to capture the level of liquidity of that asset. Furthermore $\beta_{\mathrm{LIQ},i}$ is the measure of an asset's exposure to marketwide illiquidity risk, proxied by a non-traded risk factor η_t. Likewise, $\beta_{\mathrm{MKT},i}$ captures the covariance between returns of an asset and the market. λ_{LIQ} and λ_{MKT} are the marketwide prices of exposure to liquidity and market risks, respectively.

We examine the above relation by estimating the two-step procedure defined in Equations 2.2 and 2.3 in the time series of asset returns and in the cross-section of yields, respectively:

$$R_{i,t} - R_{\mathrm{f},t} = \alpha_i + \beta_{\mathrm{MKT},i}(R_{\mathrm{MKT},t} - R_{\mathrm{f},t}) + \beta_{\mathrm{LIQ},i}\eta_t + \varepsilon_{i,t}, \quad \text{for} \quad t = 1, 2, \dots, T \quad \text{for each} \quad i; \tag{2.2}$$

$$Y_{i,t} - Y_{f,t} = \gamma_t + \kappa_t Liq_{i,t} + \hat{\beta}_{\mathrm{MKT},i}\lambda_{\mathrm{MKT},t} + \hat{\beta}_{\mathrm{LIQ},i}\lambda_{\mathrm{LIQ},t} + \nu_{i,t}, \quad \text{for} i = 1, 2, \dots, N \quad \text{for each} \quad t. \tag{2.3}$$

That is excess returns on an asset are driven by our market and liquidity factors, whereas excess yield can be explained by the level of liquidity as well as by exposure to the market and liquidity risk premiums. Note, that the liquidity beta in our model corresponds to commonality in liquidity in the Pastor and Stambaugh (2003) sense:

$$\beta_{\mathrm{LIQ},i} = \frac{Cov(\eta_t, R_{i,t})}{Var(\mu_t)} \tag{2.4}$$

The above beta captures the covariance between individual asset return and the market-wide liquidity factor. This relationship implies that the more illiquid the market is; the higher return investors would prefer.

To estimate the relationship described by Equations 2.1- 2.3, we apply the following five steps:

1. We calculate monthly asset and market returns for each asset in all three markets.

2. We define both asset and market level liquidity proxies for each of the markets at a monthly granularity and describe additional controls.

3. We construct a non-traded liquidity factor by taking the residual from an imposed AR structure of the liquidity measure in our model.

4. We describe how we proxy expected returns

5. And finally we discuss in detail the estimation strategy: given the above theory we test this relationship by means of factor models. In these models we incorporate both the level and risk aspects of liquidity. For the latter we add a non-traded liquidity factor to our analysis to see whether the risk exposure to liquidity affects prices.

2.2.1 Asset and market returns

For bond markets one can apply the standard return definition based on the ratio of consecutive prices including a correction term when coupon payments occur.[6] However, calculating returns of zero investment products is nontrivial. We define returns on inflation swaps in accordance with bond market conventions as the change in the swap rate from one period to the other multiplied by the duration of the contract. We calculate duration as that of a bond, which has a coupon rate that equals the swap's yield and maturity of the swap contract:[7][8]

$$R_{\mathrm{swap},i,t} = -(r_{i,t} - r_{i,t-1}) \cdot Dur_{\mathrm{swap},i,t} \tag{2.5}$$

In addition to asset specific returns, we also construct market returns as equally weighted average returns, where we average over all available assets at a given point in time. This is similar to for instance Amihud (2002) and Chordia et al. (2001), whereas Acharya and Pedersen (2005) also test their model on value-weighted average returns. In our sample there is no variable based on which we could weigh swap contracts and issuance amount weighted bond return figures are virtually identical to the ones that we apply. Moreover, by equally weighting our assets, we can compensate for overrepresentation of larger thus potentially more liquid bond issues.

2.2.2 Liquidity proxies and additional controls

For our empirical analysis, we need to define both individual asset and market level liquidity proxies. Unfortunately, in case of all three markets the data on the directly observable candidate, on the bid-ask spread, do not seem to be reliable over our sample period so we need to look for alternatives. Given the limited data availability in this specific market, in order to capture swap market liquidity, we construct measures that can be derived from the above swap return definition. Therefore, we propose the proportion of zero returns and the Roll measures as liquidity proxies. Similarly to Bekaert et al. (2007), the proportion of zero returns over a given period is measured as the percentage of days with zero returns over a month. This measure is particularly useful for asset classes where data availability is limited, like for the case of inflation swaps. Our second swap liquidity proxy is based on Roll (1984). The Roll measure is derived from the

[6] This correction applied at the coupon date is essentially identical to what one would do to adjust returns of a dividend paying stocks at dividend date.

[7] The minus of the quantity is taken to make the return definition resemble that of the convention for bonds.

[8] An alternative return proxy is the swap breakeven rate itself, which is a similar concept to breakeven inflation rate.

autocovariance of returns, which captures the transitory component in observed prices. It is calculated as the scaled autocovariance for the case when it is strictly negative – otherwise the measure is truncated at zero.[9]

$$Roll_{i,t} = \begin{cases} 2\sqrt{-\text{cov}(R_{i,t}, R_{i,t-1})}, & \text{if} \quad \text{cov}(R_{i,t}, R_{i,t-1}) < 0 \\ 0, & \text{otherwise.} \end{cases} \tag{2.6}$$

Measuring bond illiquidity also poses challenges, as many of the commonly applied measures, such as bid-ask spreads are unreliable, whereas the proportion of zero returns or the Roll measure are uninformative over our sample period.[10] Consequently, we turn to certain asset characteristics that are linked to a security's liquidity.

Houweling et al. (2005) propose issued amount and age of bond issues as such measures. The reasoning behind a bond's age being a proxy for liquidity is simple: the more time passes since issuance, the more likely that a bond gets locked-up in buy-and-hold investors' portfolios which decreases its liquidity. This suggests a positive relationship between illiquidity and age, whereas issued amount is negatively related to the latter: larger issues tend to be more liquid. We define age as the days since issuance, whereas we used the natural logarithm of the original issued amounts.

In the spirit of Krishnamurthy (2002), we also include an indicator variable, that equals 1 if the given issue is on-the-run – meaning that it is the latest issued security of its tenor – and zero otherwise. It can be concluded from the issuance calendars that TIPS are issued on an annual basis, whereas the cycle for nominal bonds is six months, thus the dummies are set accordingly. Based on the idea that new issues are more liquid as previous ones therefore carry smaller liquidity premium if at all, we expect the sign of this variable to be negative.

In addition to the previous liquidity proxies, we construct additional controls that we include in our analysis, such as yield volatility or a control for the slope of term structure of bonds. Yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is taken over the different maturities for a given month. This definition is the same across both swaps and bonds.[11] For bonds we also include time-to-maturity,

[9] A convenient interpretation of the Roll measure is the implied bid-ask spread, which is the bid-ask spread that can be derived from the autocovariance of returns. However, an important underlying assumption of the Roll measures interpretation is that asset returns are identically and independently distributed over time.

[10] More specifically, over our sample period all bonds are traded on a daily basis, thus the proportion of zero returns does not provide us with either cross-sectional or time series variation as for the Roll measure, the basic assumption of i.i.d. returns is not fulfilled because bond returns have positive autocorrelation in the sample.

[11] A bonds yield volatility could also serve as a proxy for liquidity, since it functions as a measure of

which is defined as the days until maturity of a given issue. This variable is supposed to control for the linear maturity structure of bonds and incorporate the slope effect of the term structure.

2.2.3 The illiquidity factor

To answer whether liquidity risk is priced, we turn to specifying marketwide liquidity proxies. These measures are calculated on a monthly frequency. By means of the aggregate volume data[12] on primary dealer bond transactions, it is a natural choice to construct the ILLIQ measure[13] of Amihud (2002). We define the measure as a ratio of weekly absolute bond market returns over weekly aggregate trading volume, where the volume is aggregated across all dealers and all securities within their class. As most of our variables are at a monthly frequency, we smooth this variable by taking its average over the four observations in a given month.

We use the ILLIQ measure to construct the illiquidity factor in our benchmark analysis, but in the robustness checks we incorporate two alternative measures. In the first case (BOND_PC), we take the first principal component of the ILLIQ measures corresponding to TIPS, all nominal Treasuries and to 10-year nominal bonds. Our second alternative measure (ALL_PC) aims to capture a wider definition of liquidity and investor sentiment, by incorporating all bond ILLIQ measures, the TED spread, alongside with the VIX index and the average Roll measure across all swap maturities for a given month. For the swap market we simply apply either the average Roll measure as defined previously or the ALL_PC measure. This principal component approach is inspired by Korajczyk and Sadka (2008).

To examine the effect of liquidity on asset prices, we construct a factor that captures marketwide liquidity risk. In unreported regressions we show that our liquidity measures

yield uncertainty. In case of more volatile yields, investors and especially market makers are uncertain of the bonds value, which increases bid-ask spreads and therefore leads to lower liquidity. On the other hand, volatility is likely to be correlated with potentially omitted factors in our models, thus we decided to include it in our regressions to partially control for those factors. Note that this makes it more difficult for our liquidity affects to survive.

[12] Volume figures on primary dealer transactions can be accessed via the Federal Reserve of New York, where they provide information on primary dealer transactions and holdings that are reported on a weekly frequency. The published figures are aggregated over all primary dealers for a given security class and week.

[13] In this study we calculate Amihuds measure based on holding period returns, while theoretically it should be based on price returns. The difference between holding period and price returns is that the former contains accrued interest and corrections for coupon payments, while the other is based on clean prices. These mechanical effects should not greatly influence the price impact, which we confirm by recalculating the measure based on price returns. The resulting ILLIQ measures are virtually identical to those based on holding period returns, with a correlation of 99.9982%.

are persistent, thus we can define this risk as the surprise or unexpected liquidity, which is the difference between expected and realized liquidity. Thus for the aforementioned market-wide liquidity measures we define the non-traded risk factor the following way:

$$\eta_t = Liq_t - \mathbb{E}_{t-1}[Liq_t] \tag{2.7}$$

To compute these innovations, we impose an autoregressive structure for the liquidity measures, similarly to Acharya and Pedersen (2005). To determine the number of lags included in these models, we require the residual or unexpected liquidity to behave as white noise. Consequently, we propose an AR(3) structure for the ILLIQ measure, and AR(1) for the average Roll measure, and for the principal components BOND_PC and ALL_PC. The underlying assumption of the above factor construction is that the AR coefficients of the series entering the principal components are identical or at least very similar. In the baseline specifications this assumption is met: the ILLIQ measures exhibit very similar time-series patterns and their AR coefficients are also quite similar: 0.58 and 0.62 for nominal Treasuries and TIPS, respectively. For swaps, we impose the AR structure on the Roll measure directly, thus for the benchmark specification of each market segments the above assumption is met.

2.2.4 Measuring expected returns

We should point out that our estimation strategy differs from the standard asset pricing approach in two aspects. First, we run our tests not on pre-sorted portfolios but on individual assets to be able to take advantage of the larger cross-sectional variation. This approach is inspired by Ang et al. (2008). Second, in asset pricing tests, one usually proxies expected returns with their realized historical counterparts. However, in our case returns seem to be too noisy, thus we turn to yields, which under a special set of assumptions can be viewed as a forward-looking proxy for expected returns. This approach is similar to Pflueger and Viceira (2015) as they also look at yields to identify liquidity premium in TIPS prices.

These assumptions vary across our three assets. In general, we have to assume that markets are frictionless and that the term structure of expected returns is flat. For nominal bonds this relationship holds if we assume that yields follow a random walk process. For TIPS, in addition to the previous assumption we need that inflation is constant in expectation and it is independently and identically distributed with yields. For swap one would ideally show that the swap rate equals the breakeven rate. In frictionless markets the breakeven inflation rate does not contain any inherent risk premiums, thus it can be proxied by the difference of nominal and real yields. Then the difference between two random walk processes of these yields would also follows random walk dynamics.

2.2.5 Estimation strategy

In this section we study how liquidity can affect expected returns. For that we estimate the marketwide premiums on market and liquidity factors as well as on our liquidity proxies.

In light of existing evidence on liquidity being priced in sovereign bond markets (Krishnamurthy (2002), Goyenko et al. (2011), Fleckenstein et al. (2014), and Pflueger and Viceira (2015) among many others), the purpose of this section is to show whether liquidity risk carries a premium in Treasury bonds. In addition, we also want to discover the aforementioned question in the context of the inflation swap market. So far no empirical evidence has been published on the relationship between inflation swaps and liquidity, despite the anecdotal evidence on the market not being perfectly liquid at all times.[14]

We approach the above question by following a two-stage Fama-MacBeth procedure in which we estimate market and liquidity betas from excess returns in the first stage, whereas in the second stage we run repeated-cross sectional regressions of yields on these betas and additional controls. We estimate betas and risk loadings in each market separately. Given the liquidity measures at hand, we are able to measure the covariation of a security's return with the market liquidity, as shown in Equation 2.4. This covariance suggests that market liquidity affect required returns positively, such that the more illiquid a market is, the higher returns investors expect which decreases the asset's price. In the first stage we run the following time series OLS regressions to obtain the betas:

$$\begin{aligned} R_{i,t} - R_{\mathrm{f},t} &= \alpha_i + \beta_{\mathrm{MKT},i}(R_{\mathrm{MKT},t} - R_{\mathrm{f},t}) + \beta_{\mathrm{LIQ},i}\eta_t + \varepsilon_{i,t}, \\ &\text{for} \quad t = 1, 2, \ldots, T \quad \text{for each} \quad i; \end{aligned} \tag{2.8}$$

where we include excess market returns and unexpected liquidity, which is the residual from the AR process discussed above. In the second stage we run repeated cross-sectional regressions of yields on the betas estimated in the previous step, asset level liquidity proxies and additional controls, such as the volatility of yields. Estimates from the repeated regressions are averages across time and the errors include both a 12-lag Newey-

[14] See for instance Fleming and Sporn (2012).

West correction[15] and account for the averaging of the coefficients:[16]

$$Y_{i,t} - Y_{\mathrm{f},t} = \gamma_t + \kappa_t Liq_{i,t} + \hat{\beta}_{\mathrm{MKT},i}\lambda_{\mathrm{MKT},t} + \hat{\beta}_{\mathrm{LIQ},i}\lambda_{\mathrm{LIQ},t} + \nu_{i,t}, \quad \text{for} \quad i = 1, 2, \ldots, N \quad \text{for each} \quad t. \tag{2.9}$$

As a result of this step we get estimates of the market price of liquidity and liquidity risk as well as each asset's individual exposure to this risk. These models can be formulated based on two opposed propositions: either we assume that these markets are perfectly segmented and all forms of liquidity are priced separately on the three markets or we price liquidity risk in fully integrated markets. The difference between the two approaches is stemming from the consequent definition of the market. In the integrated case the market return is that equally weighted average of all asset that are in positive net supply: thus nominal and indexed bonds. Given the evidence that nominal bonds are the most liquid among our test assets, this method is likely to produce larger liquidity effects for swaps and TIPS. On the other hand, in segmented markets, the average of market betas is one, whereas that of liquidity betas is zero by construction. As opposed to this, in the integrated case, our estimates are not bounded by the mathematics of factor models. We choose the segmented case to be our benchmark as these estimates are more conservative for the aforementioned reasons.

2.3 The data and the three markets

In this section, we describe in detail the data and the markets with evidence on liquidity issues for both nominal and indexed Treasuries and inflation swaps.

2.3.1 The data

The data consist of daily closing mid prices of TIPS and nominal Treasury bonds alongside with zero coupon inflation swap quotes. These data are obtained from Bloomberg and are similar to Fleckenstein et al. (2014)[17] as they span most of the existing TIPS issues and only include a fraction of the long-term nominal Treasury market. The data

[15] Later on, we are planning to incorporate a Shanken-type error correction that take into account the errors-invariable problem, which is stemming from the fact that both our liquidity measures and the consecutive betas are pre-estimated. We might incorporate clustered errors too, where we would cluster by assets. See Petersen (2009).

[16] For the exact formula see pp. 229 in Cochrane (2005).

[17] Our data are an extended and updated version of Fleckenstein et al. (2014) as we include a larger cross-section of bond issues, as well as we have a longer time span. Moreover, we complement this data with input for liquidity proxies and additional controls as described below.

contain maturity-matched[18] indexed and nominal issues[19] , whose maturities range between 2007 and 2041. The daily closing bond prices are adjusted by accrued interest following the market convention. Moreover, our sample contains inflation swap quotes that are the constant rate on a fixed contract's leg. Following Fleckenstein et al. (2014), we choose contracts with maturities ranging between 1 to 10, 12, 15, 20, 25 and 30 years. We apply the simplest approach to get intermediate (non-traded) maturities: we use a linear interpolation technique and include no correction for potential seasonal patterns in inflation[20] . We collect data for the US market from July 2004 to December 2011.

As our main purpose is to investigate the effect of liquidity on these three markets, we need to construct proxies for liquidity. Therefore, we also gather information on the bonds' issue and maturity dates, the amount issued and their coupons. To formally test whether liquidity risk is priced we download additional controls from Bloomberg, such as the TED spread and the VIX index next to deriving measures from prices themselves. We complement these data with zero coupon yield curves that Datastream constructs from the par yield curve.

To construct our benchmark liquidity proxy, we need volume figures on primary dealer transactions. These data can be accessed via the Primary Dealer Historical Search database of the Federal Reserve of New York, where they have information on primary dealer transactions and holdings that are reported on a weekly frequency. The published figures are aggregated over all primary dealers for a given security class and week.

Finally, to run the asset pricing tests we also obtain risk free rates from Kenneth French's website and risk free yield from St. Louis Fed's FRED database.

2.3.2 Constituent asset markets

In this section we provide a short description of the TIPS and inflation swap markets specifically focusing on market characteristics that could lead to illiquidity. We also contrast the liquidity features of the three markets under scrutiny based on prior empirical work.

[18]For the exact procedure of maturity matching, see Fleckenstein et al. (2014).

[19]The original sample consisted of 41 TIPS and 40 maturity-matched nominal issues. However, for the asset pricing test we decided to apply two filters: we omitted issues from the sample that had less than 24 months of data and also only kept observations up to six months before a bonds maturity. Interestingly both bonds after issuance and quotes preceding maturity are considerably more volatile than the rest of the sample and they resulted in extreme beta and premiums estimates.

[20]The inflation index based on which both the principal amount of TIPS and swap contracts are adjusted on a daily basis is CPI-U or CPI for All Urban Consumers.

The TIPS market

The first TIPS auction took place in 1997 and ever since the market gradually grew into one of the largest and most-actively traded fixed income markets in the world (Fleckenstein et al., 2014). As of the end of our sample period 41 individual TIPS issues have been auctioned on a regular cycle, with five-year, 10-year and 30-year maturities.[21]

TIPS in most respects are similar to nominal Treasuries, the main difference being that the principal amount is adjusted on a daily basis to changes in CPI to All Urban Consumers.[22] This implies that semiannual coupons, that are a fixed percentage of the principal linked to changes in inflation, also vary over time. Another noteworthy feature of TIPS is the embedded deflation option, which protects investors from losses: in any case investors are entitled to the maximum of the final principal amount or its inflation-adjusted counterpart. Despite the growing size of the market, an increasing number of studies have shown that TIPS carry a liquidity premium compared to their nominal counterparts of similar maturities (Fleckenstein et al., 2014; Pflueger and Viceira, 2015; Haubrich et al., 2012; Campbell et al., 2009). Moreover, we also know from those studies that liquidity carries a positive premium, as one would expect in a positive net supply market.

The inflation swap market

Kerkhof (2005) argues that the US zero-coupon inflation swap market has been a rapidly growing segment of the inflation derivatives market in the past decade as market participants began making markets to hedge their inflation risk exposures. However, the current size of the market is still about a couple of percent that of nominal interest rate swaps and is atomic compared to Treasury securities. In line with this, Fleming and Krishnan (2012) report that there are relatively few trades occurring in this market.

An inflation swap is a bilateral derivative transaction in which one party agrees to swap a fixed payment to a floating one that is tied to inflation, for a given notional amount and period of time. Inflation swaps, similarly to TIPS, are also linked to CPI-U, and the fixed rate is negotiated in over-the-counter transactions that are traded in a dealer-based market Fleming and Krishnan (2012). The most frequently traded inflation swap contracts are the zero coupon contracts, in which cash flows are only exchanged at the maturity of the contract.

[21] Previously TIPS with 20-year maturities were also issued by the Treasury.

[22] This is a non-seasonally adjusted inflation rate corresponding to urban consumers in the US.

So far, only a handful of studies investigated the breakeven rate or its relationship with inflation swaps (Campbell et al., 2009; Gürkaynak et al., 2010; Christensen and Gillan, 2011; Pflueger and Viceira, 2015; Fleckenstein et al., 2014), but as Fleming and Krishnan (2012) point out empirical evidence on inflation swap liquidity is still lacking: no study ever considered modeling the liquidity of this market. Despite that size of the market and the fact that trades occur rarely suggest that liquidity is likely to have an effect on swap returns, we cannot predict its expected direction. Bongaerts et al. (2011) show that a battery of factors, such as non-traded risk exposures in investors' portfolios, individual risk aversion liquidity's correlation with investors' hedging demands, determines the direction of liquidity's effect on markets that are in zero net supply (e.g.: derivative markets). Since these factors are unobservable, especially on the aggregate level, we cannot predict the expected sign of liquidity.

The nominal Treasury market and liquidity

In fact, many claim the nominal Treasury bond market to be the most liquid and most frequently traded fixed-income market in the world and thus it is often taken as a reference point in investigating other securities' liquidity. Krishnamurthy (2002) uses the commercial paper-T-bill spread to capture changes in liquidity demand, whereas Longstaff (2004) applies the Refcorp-Treasury spread to capture flight-to-liquidity premium in economically distressed times. Pflueger and Viceira (2015) treat nominal Treasuries as perfectly liquid to quantify the premium inherent in TIPS returns and determine bond return predictability.

On the other hand, Krishnamurthy (2002) has shown that the liquidity of nominal bonds does vary significantly over the issuance cycle therefore liquidity premium can indeed be found in Treasury returns too. For this reason, we also take a look at these bonds' liquidity.

Note that although our sample contains all available TIPS that are issued prior to December 30, 2011, it could contain significantly more nominal issues. This is because the data has been collected in the spirit of Fleckenstein et al. (2014), such that it comprises of maturity-matched indexed and nominal bond pairs alongside with inflation swap contracts. Besides, we believe that our sample of bonds exhibits enough variation in liquidity features such that our results can be generalized to the entire population of long-maturity nominal Treasuries. In this case having a larger sample of bonds would only make our results stronger.

2.4 Empirical results

This section presents the result of estimating the two-stage model described in Section I. We first discuss the descriptives and our estimated betas from Equation 2.4, alongside with the properties of the liquidity factor. We proceed with showing our benchmark results for all three markets. Next, we also provide robustness tests including other liquidity measures or alternative assumptions regarding the relationship of our markets. And finally, we demonstrate how our results apply to the trading strategy described in Fleckenstein et al. (2014).

2.4.1 Descriptives, betas and the illiquidity factor

Table 2.1 contains the descriptives of our liquidity proxies for all three markets in our sample, whereas Table 2.2 provides the distribution of the betas estimated in the first stage of our analysis. Table 2.1 reports all quantities but the ILLIQ measure and the on-the-run dummy in percentages and shows the main characteristics and the distribution of our liquidity measures. In the swap market the average yield is 2.47%. The relative yield volatility measure by construction equals to zero, but individual issues can significantly differ from the cross-sectional average. The Roll measure implies an average bid-ask spread of 24.5 basis points. On average 5.82% of the times we have zero returns on this market this suggests no trading activity on average 1.8 days a month.

For TIPS the average yield is 74 basis points and yield volatility of individual issues varies in a wider range than for swaps. The age of the average bond in our sample is 4.38 years with average time to maturity of 9.12 years. The average issue size is $16.1 billion. The dummy variable shows that 12.67% of the issues are on-the-run. We also present the ILLIQ measure, a price impact proxy in our sample.[23]

In comparison, nominal Treasuries have higher yields, on average 1.58%, with lower relative yield volatility than TIPS or inflation swaps. These bonds are older, with the average age of 5.86 years with also somewhat longer time to maturity, 9.48 years. The average nominal issue is also larger than that of TIPS, with $21 billion. Only 4% of the issues are on the run, which is less than for TIPS.[24]

Figure 2.1 depicts the time evolution of our non-traded liquidity factors that are residuals from autoregressive processes: AR(3) for ILLIQ measures concerning TIPS, and AR(1) for

[23]The ILLIQ measure is the absolute dollar change triggered by volume, however this number depends on rescaling.

[24]This is not surprising given that the nominal bond issuance cycle is half that of TIPS, thus 6 months.

the nominal Treasury ILLIQ and the average Roll measure. The TIPS and nominal bond series are relatively highly correlated, with a correlation coefficient of 0.6131, whereas their correlation with the inflation swap market series is 0.1680 and 0.4133, respectively. Apparently the TIPS liquidity factor has larger swings and more spikes than the other two series, whereas the nominal bond and inflation swap factors shoot up during the recent financial crisis. Both time paths are in line with anecdotal evidence and previous empirical findings of illiquid periods in the corresponding markets.

For the sake of brevity,[25] Table 2.2 focuses on the distribution of betas that we estimate from time-series regressions[26] of returns on the market excess return and the illiquidity factor. Note that our market factor in this context is practically an interest rate risk or duration factor, which explains the patters in Figures 2.2- 2.4. These graphs depict the betas sorted on average age of an issue for TIPS and nominal Treasuries and contract maturity for inflation swaps. We present result for both cases of integrated and segmented markets[27].

In the case of inflation swaps, when considered as a segmented market, we see that liquidity betas have a larger spread than market betas. Loadings on the market factor are all positive. In the integrated case, where we take market as the sum of the nominal and indexed bonds, our estimates change: the average market beta is still close to zero however certain issues load on the market factor with negative sign. We also see that the magnitude and the spread of liquidity betas substantially increase, whereas they always have a negative sign. The two panels of Figure 2.2 confirm these findings.

Figures 2.3 and 2.4 expose that in both the segmented and integrated cases TIPS and nominal bonds have strictly positive market betas, which vary in a narrower range than those of swaps. We observe a similar difference in range for segmented liquidity betas of TIPS. Ex ante we would expect TIPS prices to decrease if liquidity decreases. In contrast with expectations, when we assume markets are integrated, most TIPS issues load positively on our illiquidity factor, whereas nominal bonds tend to have negative and sizeable illiquidity betas.

[25]After applying our data filters, we estimate market and liquidity betas for 15 swap, 31 TIPS and 32 nominal bonds. Thus given the large number of individual assets, in the table we focus on the distribution of our estimates.

[26]We conclude from unreported regressions, that market betas are highly significant in case of most assets, whereas the statistical significance of individual liquidity betas varies a lot.

[27]In segmented markets, the average of market betas is one, whereas that of liquidity betas is zero by construction. As opposed to this, in the integrated case, our estimates are not bounded by the mathematics of factor models.

2.4.2 Benchmark results

We estimate our benchmark models (i) under the assumption of market segmentation; (ii) using illiquidity factors derived from ILLIQ for bonds and average Roll measure for inflation swaps; (iii) for the period between July 2004 and December 2011. To define the baseline specification, we pick models that are significant yet parsimonious. Consequently, for nominal Treasuries and inflation swaps we pick the model with the market and the illiquidity factors as our baseline specifications; whereas for TIPS, based on unreported univariate regressions, we extend the latter model with two characteristics: age and issued amount of a bond. We also include the economic effect of risk premiums in the last rows of the tables, which are based on the monthly holding periods implied by the regressions. To capture the impact of market and liquidity risks, we calculate the interquartile spread: the estimated price of risk multiplied by the difference between the betas corresponding to the first and third quartile in the cross-section of betas. Parameter estimates of the benchmark cases can be found in the first columns of Tables 2.3, 2.5, and 2.7.

The first column of Table 2.3 presents the benchmark case for inflation swaps. Although there is no prior literature on inflation swap liquidity, anecdotal evidence from Fleming and Krishnan (2012) suggest it may have an effect on swap yields. However, we cannot predict its expected direction as there are many factors that determine how liquidity impacts markets that are in zero net supply.[28] The key result that we find for inflation swaps is that both market and liquidity risks are priced on this market. The market price of liquidity risk is positive and it is statistically significant, nevertheless the implied monthly economic impact is 1.65 basis points. This effect is small as is often the case in derivative markets. On the other hand, the economic impact of market risk is a sizable 43.92 basis points.

Table 2.5 reports our benchmark case for TIPS. The growing literature on TIPS illiquidity suggests TIPS prices to convey liquidity discount.[29] Therefore, we expect liquidity risk to be priced. As for the included characteristics, for age we expect a positive sign as the more time passes since issuance, the more likely an issue gets locked up in buy-and-hold investor portfolios, which increases illiquidity. On the other hand, larger bond issued tend to be more liquid, therefore the expected sign of size of an issue is negative. Our main finding is that for TIPS the effect of illiquidity risk is dominated by asset characteristics such as age and the size of an issue. While market risk is priced, illiquidity is both statistically and economically insignificant. Age of an issue is both statistically and economically

[28]Bongaerts et al. (2011) show that non-traded risk exposures in investors portfolios, individual risk aversion liquiditys correlation with investors hedging demands are such factors. Since these are unobservable especially on the aggregate level, we cannot predict the expected sign of liquidity.

[29]Including Campbell et al. (2009), Christensen and Gillan (2011), Pflueger and Viceira (2011, 2015), Haubrich et al. (2012), Fleckenstein et al. (2014)

important driver of TIPS liquidity, with an impact of 41.71 basis points. If a bond gets one year older, its yield will increase with 9.09 basis points, which implied a decrease in its price. Moreover, the size of an issue also matters its effect is -9.03 basis points. Similarly to age, once a TIPS issue gets 1% larger, we expect its yield to decrease by 0.31 basis points, thus its price increases.

Turning to nominal Treasury notes and bonds, the benchmark case can be found in the first column of Table 2.7. In line with previous literature, we presume nominal bonds are more liquid than other securities – as in Krishnamurthy (2002), Longstaff (2004) or Pflueger and Viceira (2015), among many others. Our benchmark specification focuses on the two factors: market and illiquidity. Ex ante, the sign of the illiquidity premium is not clear, as for instance Fontaine and Garcia (2012) find negative liquidity premium in nominal Treasuries which makes these securities good liquidity hedge in periods of flight-to-liquidity. We find that in the nominal Treasury market illiquidity risk is priced and carries a positive but fairly small premium of 13.13 basis points. At the same time, the economic effect of the market is substantial, with 183.22 basis points.

2.4.3 Robustness tests

To check the robustness of our benchmark specifications, we include additional controls, as well as test the effect of the assumption of integration. In unreported results we also construct and assess other liquidity factors alongside with splitting our period into subsamples.[30] We also test whether the price of liquidity risk is different in different liquidity regimes: when we restrict our sample to those months when our aggregate market illiquidity factor increases, and separately to those when it increases.

Taking another look at Table 2.3, we find that the fore mentioned benchmark case is robust to the inclusion of asset characteristics and controls, such as the proportion of zero returns or the volatility of swap yields. In all cases the economic impacts of market and liquidity risk exposures do not change substantially either in sign or magnitude. In contrast, we also perform a similar analysis under the assumption swap and bond market integration. From Table 2.4 we see that both the market and illiquidity betas are priced. These effects are also highly significant and robust to the inclusion of volatility. It is in line with our expectations that the sign corresponding to the price of illiquidity changes and the implied premium also doubles in size: it is between 18.2 and 34.1 basis points. The negative premium implies that the less liquid a swap issue is, the lower the price and thus the higher the expected return on and the yield of that asset is.

[30] We are planning to incorporate these analyses in later versions.

For TIPS we observe that age and the size of an issue matter more than liquidity risk. Other columns in Table 2.5 show accordingly. However, if we only include the market and the illiquidity factor, in column 2, liquidity risk seems to carry a significant and sizeable premium of -22.6 basis points, and this remains so after the inclusion of the on-the-run dummy, issued amount and yield volatility. Nevertheless, once age or time-to-maturity is included, these variables wipe out the factor's significance. Similarly to swaps, we repeat the analysis for integrated markets – Table 2.6 contains the corresponding results. In general results do not change: the magnitude of the effects is both statistically and economically similar to the previous case. Therefore, we conclude this market is not as sensitive to this assumption as inflation swaps.

Looking at nominal Treasuries, we show that the market and illiquidity premiums are robust to inclusion of asset level characteristics and controls. Interestingly age next to being significant in all specifications, has the wrong sign.[31] Our time-to-maturity control suggests the slope of the term structure to matter, besides other variables, such as the on-the-run dummy, issued amount and yield volatility, are never significant. In comparison if we take the integrated market case in Table 2.8, the characteristics and controls seem to carry a more important role. The market and illiquidity factors are highly significant with similar premiums estimates as under market segmentation assumption. Age still has the wrong sign, but now the size and yield volatility of an issue are significant determinants of yields.

In Table 2.9, we present results conditional on whether market liquidity is increasing or decreasing. To do so, we pool together months where changes of market liquidity are either positive or negative, respectively. What we find is that in all three market segments both the signs, as well as the magnitudes of the estimated risk premia are stable across liquidity regimes. This suggests that our results are robust to liquidity regimes, however, we could further extend the analysis to examine if Treasury and inflation swap markets react to extreme downside liquidity Ruenzi et al. (2016).

2.5 The relative pricing of TIPS and nominal Treasuries

In this section, we apply the result of the previous section to look at the effect of liquidity on the relative pricing of nominal and indexed Treasuries that we inspect by the trading

[31] We suspect that age might have a nonlinear relationship with yields; therefore, in the future we are planning to examine such effects.

strategy of Fleckenstein et al. (2014).[32] The idea behind their TIPS-Treasury arbitrage is simple: an investor matches the maturities and payoffs of a nominal bond issue and its synthetic counterpart. The latter is essentially an inflation swapped-indexed bond, whose cash flows are converted to fix payments exactly matching that of the corresponding nominal bond.

To replicate this strategy, an investor should buy a TIPS issue and short a nominal bond at the same time. Additionally, she needs to execute a zero-coupon inflation swap contract with the same maturity and notional amount as the TIPS coupon – and repeat this for each coupon and for the principal amount, which results in the execution of an entire swap portfolio. The rationale for swapping the bond is that the sum of the two cash flows is constant and equal to the nominal coupon or principal. The investor also needs to take a small position in Treasury STRIPS[33] due to the disparity in the nominal and TIPS coupon payments. Based on this logic the investor applies these steps to all coupon payments, which results in the successful conversion of the TIPS variable cash flow stream to the fixed one of the corresponding nominal bond.

In sum, the investor would short sell the nominal bond, buy the TIPS issue and hold portfolios of zero-coupon inflation swap contracts and Treasury principal STRIPS. The latter three components exactly replicate the fixed periodic coupons and the principal of the nominal bond. Finally, to calculate what Fleckenstein et al. (2014) call mispricing, we first price the synthetic bond by calculating the yield to maturity and then the price of the replicating portfolio.[34] Thus, if the resulting prices of the nominal bond and the replicating portfolio differ, a potential arbitrage opportunity arises.

To incorporate the direct effect of liquidity, we adjust the yields of nominal and indexed Treasuries by the estimated premiums of the benchmark cases. We also include the effect of liquidity on inflation swap positions concerning every coupon payment within the strategy. To take account of the liquidity-adjustment in swap contracts[35], we calculate

[32]A minor change we apply to Fleckenstein et al. (2014) is that when we calculate the mispricing, instead of the ones with accrued interest, we apply clean prices. The reason is that the coupon date of most bond issues on our sample coincide, therefore the price differential based on dirty prices has a jagged pattern, representing the average accrued interest at a given point in time. To circumvent this problem, we use clean prices.

[33]We are aware that STRIPS might also be exposed to liquidity issues (see for instance Daves and Ehrhardt (1993) or Jordan et al. (2000)), however Bühler and Vonhoff (2011) find that principal STRIPS are less affected. As the trading strategy presented in this paper uses principal STRIPS, we are less concerned that small positions taken in these assets would carry a sizable liquidity premium that could distort our current results. A potentially negative premium, which is in line with previous findings, would only work in our favor by reducing the mispricing.

[34]Practically we also need to adjust the price of the nominal bond for the potential maturity mismatch between the two securities.

[35]As the strategy consists of zero coupon inflation swaps, we only need to apply this formula once, corresponding to the maturity o the underlying swap contract.

the difference between the fixed and the floating cash flows of the inflation swap contract by the following formula:

$$Value_{\text{swap},t} = n \cdot \hat{\beta}_{\text{LIQ,swap}} \lambda_{\text{LIQ}} \frac{s(1 + y_{\text{swap},n,t})^n}{(1 + y_{\text{zc},n,t})^n}. \tag{2.10}$$

That is the value of the liquidity corrected position is the estimated price of liquidity for a given swap contract, $\hat{\beta}_{\text{LIQ,swap}} \lambda_{\text{LIQ}}$ multiplied by the swapped TIPS coupon (s) discounted by the appropriate rate from a nominal zero-coupon yield curve, where $(1 + y_{\text{swap},n,t})^n$ is the forward of the contract with n years maturity[36] and $(1 + y_{\text{zc},n,t})^n$ is that of a zero-coupon bond with the same maturity. $y_{\text{swap},n,t}$ is the quoted swap yield of an n-maturity contract at time t, whereas $y_{\text{zc},n,t}$ is the nominal zero-coupon yield of the same maturity at the same point in time.

The result of the liquidity correction of the price differential can be found in Figure 2.5 and Table 2.10. Figure 2.5 compares the replicated 'mispricing' of Fleckenstein et al. (2014)) and its adjusted counterparts under the assumptions of segmented and integrated markets. Our key result is that once we take out the estimated liquidity premiums from prices, the price differential shrinks considerably, if not disappears. This is in accordance with our expectations and the results from previous sections. The shrinkage of price difference is true for both specifications, although the effect of our liquidity adjustment is larger for the segmented market case. Table 2.10 confirms these findings: whereas all values regarding the replicated series are positive, the corrected series based on market segmentation are mostly negative. We also define the difference between the mispricing series, as the difference between the original strategy and our corrected version. We find that this disparity is always positive and often times considerable in magnitude, especially in proportion to the uncorrected series.

2.6 Conclusion

We show that in both index-linked bond markets and inflation swap markets liquidity is an important determinant of prices. We do so by estimating a model with both a liquidity risk factor and asset-specific liquidity characteristics. To estimate the effect of liquidity risk, we measure an asset's exposure to our non-traded liquidity factor. In addition to this liquidity risk exposure, the level of liquidity is proxied by asset-level characteristics. We also study liquidity effects in nominal bonds in a similar way, so that in total we analyze liquidity premiums in three markets. We conduct our analyses based on two alternative

[36]For non-traded and fractional maturities we apply linear interpolation to get the forward rate, as well as the liquidity premium.

assumptions – we either propose the three markets being segmented, such that prices are independently determined, or integrated markets.

Additionally, we also scrutinize whether the exposure to liquidity and liquidity risk could explain the persistent difference in relative bond prices, as documented in Fleckenstein et al. (2014). They show that there exist substantial price differences between a nominal Treasury bond and its synthetic counterpart – a swapped TIPS issue. We provide evidence that when controlling for liquidity, a large part of this apparent mispricing disappears.

Yet on the empirical side several extensions of the paper are possible and considered. In important question to be addressed is whether our liquidity proxies are affected by unconventional monetary policy actions, like quantitative easing during the crisis, as studied by Christensen and Gillan (2013) or D'Amico and King (2013). Some other issues also remain to be solved. For instance, we are planning to add a stylized model that justifies our use of yields as a proxy of expected returns. We also want to incorporate more robustness checks regarding other liquidity factors and dividing our data into subsamples where premiums are separately estimates. One could also consider incorporating time-varying risk premiums.

Table 2.1
Descriptive statistics: swap and bond markets

The table presents descriptive statistics for variables used in the two-stage estimation. Panel A present variables for the analysis of inflation swap markets, whereas Panel B and C show those for TIPS and nominal Treasuries, respectively. Swap yields are percentage quoted rates of a swap contracts, whereas yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields. The Roll measure is the scaled autocovariance of inflation swap returns, while the proportion of zero returns is measured as the percentage of days with zero returns over a month. Age and time-to-maturity are defined relative to the issue and maturity dates, respectively. The on-the-run dummy is an indicator variable, that equals 1 if the given issue is the latest issued security of its tenor and zero otherwise. ILLIQ is the monthly average ratio of weekly absolute bond market returns over weekly aggregate trading volume. Yields, volatilities, the Roll and the zero returns measures are in percentages, age and time-to-maturity are measured in years. The data correspond to the sample period between July 2004 and December 2011.

Panel A: Descriptive statistics of inflation swap markets

	Mean	St. Dev.	Min	p25	p75	Max
Swap yield	2.47	0.71	-3.83	2.33	2.87	3.41
Yield volatility	0	0.06	-0.35	-0.02	0.01	1.14
Roll measure	0.25	0.43	0	0	0.31	5.58
Proportion of zeros	5.82	13.63	0	0	5	100
Average Roll m.	0.25	0.24	0	0.09	0.36	1.92

Panel B: Descriptive statistics of TIPS

	Mean	St. Dev.	Min	p25	p75	Max
TIPS yield	0.07	1.67	-3.02	-1.26	1.34	7.47
Age	4.38	3	0	1.95	6.54	13.72
Time-to-maturity	9.12	6.93	16	3.80	15.48	27.73
Issued amount	23.50	0.39	22.34	23.43	23.72	24.06
Yield volatility	0	0.07	-0.42	-0.02	0.01	1.08
On-the-run	0.13	0.33	0	0	0	1
ILLIQ	1.51	0.45	0.87	1.19	1.77	2.91

Panel C: Descriptive statistics of nominal Treasuries

	Mean	St. Dev.	Min	p25	p75	Max
Nominal yield	1.58	1.31	-0.67	0.46	2.59	4.51
Age	5.86	3.98	0	2.50	8.63	16.88
Time-to-maturity	9.48	7.10	0.13	3.90	16.56	26.56
Issued amount	23.77	0.42	23.07	23.43	24.06	24.92
Yield volatility	0	0.03	-0.17	-0.01	0.01	0.23
On-the-run	0.05	0.21	0	0	0	1
ILLIQ	2.38	0.55	1.55	2.01	2.56	5.22

Table 2.2
Beta estimates

The table presents descriptive statistics for betas estimated from the time-series regression of asset returns on market and non-traded illiquidity factors. Panel A present variables for the analysis of inflation swap markets, whereas Panel B and C show those for TIPS and nominal Treasuries, respectively. We estimate market and illiquidity betas for 15 swaps, 31 TIPS and 32 nominal Treasury issues in our sample that spans the period between July 2004 and December 2011.

Panel A: Inflation swap market

	Mean	St. Dev.	Min	p25	p75	Max
Segmented market β	1	0.57	0.22	0.57	1.40	2.18
Segmented illiquidity β	0	1.50	-4.14	-0.29	0.68	2.95
Integrated market β	0.02	0.26	-0.43	-0.27	0.24	0.43
Integrated illiquidity β	-4.09	3.09	-12.78	-5.59	-2.37	-0.30

Panel B: TIPS market

	Mean	St. Dev.	Min	p25	p75	Max
Segmented market β	0.96	0.45	0.29	0.53	1.42	1.92
Segmented illiquidity β	0.01	0.39	-0.90	-0.26	0.35	0.47
Integrated market β	0.99	0.59	0.16	0.46	1.61	2.04
Integrated illiquidity β	0.46	0.44	-0.68	0.06	0.73	1.17

Panel C: Nominal Treasury market

	Mean	St. Dev.	Min	p25	p75	Max
Segmented market β	0.96	0.57	0.17	0.39	1.66	1.90
Segmented illiquidity β	-0.07	0.25	-0.67	-0.16	0.10	0.51
Integrated market β	0.95	0.56	0.19	0.37	1.66	1.85
Integrated illiquidity β	-0.37	0.37	-1.21	-0.65	-0.12	0.50

Table 2.3
Monthly swap yields and illiquidity – Market segmentation

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of markets being segmented. The dependent variable is the inflation swap yield. Market and illiquidity betas are estimated based on Equation 2.4 as loadings on the market and non-traded illiquidity factors. The Roll measure is calculated as the scaled autocovariance of returns for the case when it is strictly negative otherwise the measure is truncated at zero. The proportion of zero returns is measured as the percentage of days with zero returns over a month, whereas (lagged) yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is take over the different maturities for a given month. Displayed coefficients are average figures from monthly repeated cross-sectional regressions, where errors take into account the averaging and include a 12-lag Newey-West correction. The economic impact is captured by the interquartile spread: we multiply the coefficient by the difference between the betas that correspond to the first and third quartile in the cross-section of betas. The sample period is July 2004 until December 2011. t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	Benchmark	(2)	(3)	(4)	(5)	(6)
Market beta	0.5338	0.4269	0.5796	0.5330	0.4518	0.4466
	(3.91)***	(4.38)***	(3.25)***	(3.77)***	(3.93)***	(3.80)***
Illiquidity beta	0.0171	0.0092	0.0048	0.0226	0.0056	0.0106
	(2.18)**	(2.21)**	(1.06)	(1.91)*	(1.67)*	(1.96)*
Yield volatility$_{t-1}$		-2.0809			-1.7160	-1.5542
		(1.78)*			(-1.62)	(-1.64)
Roll measure			-0.1647		0.0516	0.0588
			(-0.86)		(0.65)	(0.79)
Proportion of zero returns				-0.0158		-0.0091
				(-1.46)		(1.67)*
Intercept	1.9402	2.0428	1.9189	1.9602	2.0200	2.0401
	(6.88)***	(8.63)***	(6.57)***	(7.03)***	(8.18)***	(8.27)***
Adj. R^2	0.67	0.85	0.71	0.68	0.86	0.87
Number of obs.	1,350	1,335	1,350	1,350	1,335	1,335
Impact of market risk	0.4392	0.3512	0.4768	0.4385	0.3717	0.3674
Impact of liquidity risk	0.0165	0.0089	0.0046	0.0218	0.0053	0.0102

Table 2.4
Monthly swap yields and illiquidity Market integration

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of integrated markets. The dependent variable is the inflation swap yield. Market and illiquidity betas are estimated based on Equation 2.4 as loadings on the market and non-traded illiquidity factors. The Roll measure is calculated as the scaled autocovariance of returns for the case when it is strictly negative otherwise the measure is truncated at zero. The proportion of zero returns is measured as the percentage of days with zero returns over a month, whereas (lagged) yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is take over the different maturities for a given month. Displayed coefficients are average figures from monthly repeated cross-sectional regressions where errors take into account the averaging and include a 12-lag Newey-West correction. The economic impact is captured by the interquartile spread: we multiply the coefficient by the difference between the betas that correspond to the first and third quartile in the cross-section of betas. The sample period is July 2004 until December 2011. t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	(1)	(2)	(3)	(4)	(5)	(6)
Market beta	0.4531	0.2626	0.4130	0.4729	0.2508	0.2674
	(2.44)**	(2.38)**	(2.38)**	(2.36)**	(2.40)**	(2.33)**
Illiquidity beta	-0.0662	-0.0462	-0.0621	-0.0607	-0.0402	-0.0353
	(5.03)***	(6.25)***	(4.23)***	(5.34)***	(5.39)***	(5.92)***
Yield volatility$_{t-1}$		-3.2833			-3.3298	-2.8924
		(2.65)***			(2.60)**	(2.82)***
Roll measure			-0.1200		0.3478	0.3463
			(-0.64)		(1.22)	(1.20)
Proportion of zero returns				-0.0051		-0.0084
				(-0.62)		(-1.40)
Intercept	2.1947	2.2759	2.1716	2.2202	2.2578	2.2978
	(10.85)***	(13.04)***	(10.41)***	(12.03)***	(12.52)***	(13.43)***
Adj. R^2	0.59	0.75	0.62	0.61	0.78	0.80
Number of obs.	1,350	1,335	1,350	1,350	1,335	1,335
Impact of market risk	0.2330	0.1351	0.2124	0.2432	0.1290	0.1375
Impact of liquidity risk	-0.0341	-0.0238	-0.0319	-0.0312	-0.0207	-0.0182

Table 2.5
Monthly TIPS yields and illiquidity – Market segmentation

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of markets being segmented. This implies that the market return is the equally weighted average of indexed Treasuries. The dependent variable is the TIPS yield. Market and illiquidity betas are estimated based on Equation 2.4 as loadings on the market and the non-traded illiquidity factors. We define age as the years passed since issuance, and time-to-maturity as the years until maturity. The on-the-run variable is a dummy that equals one of an issue is the newest of its tenor and zero otherwise. We use the natural logarithm of the original issued amounts; whereas (lagged) yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is take over the different maturities for a given month. The economic impact is captured by the interquartile spread: we multiply the coefficient by the difference between the betas that correspond to the first and third quartile in the cross-section of betas. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions where errors take into account the averaging and include a 12-lag Newey-West correction. The sample period is July 2004 until December 2011. t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	Benchmark	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Market beta	1.0317	0.8997	1.0488	-1.2651	0.9483	0.8976	1.3132	1.0455	1.1490	1.0608	1.0691	0.0949
	(3.65)***	(2.94)***	(3.78)***	(4.04)***	(3.18)***	(2.92)***	(4.73)***	(3.74)***	(5.19)***	(5.10)***	(5.10)***	(0.81)
Illiquidity beta	-0.0141	-0.3652	-0.0261	0.0310	-0.3053	-0.2975	-0.0602	-0.0566	0.0561	0.0434	0.0326	0.2407
	(-0.18)	(3.73)***	(-0.34)	(0.25)	(3.07)***	(3.16)***	(-0.74)	(-0.76)	(-1.08)	(-0.70)	(-0.52)	(2.95)***
Age	0.0909		0.1000					0.0931	0.0949	0.0887	0.0847	0.0686
	(9.18)***		(10.30)***					(5.82)***	(11.73)***	(10.19)***	(7.38)***	(7.46)***
Time-to-maturity				0.1573								0.0832
				(9.46)***								(4.30)***
On-the-run indicator					-0.2737			0.0019			0.0031	-0.0137
					(3.62)***			(0.03)			(0.07)	(0.70)
Issued amount	-0.3135					-0.4330				-0.3121	-0.3041	-0.2177
	(7.19)***					(10.42)***				(6.79)***	(6.40)***	(9.87)***
Yield volatilityt-1							7.1119		2.1668	0.9783	0.9843	2.7513
							(3.49)***		(1.35)	(0.58)	(0.57)	(1.98)*
Intercept	5.8625	-0.9874	-1.5563	-0.3890	-0.9789	9.1904	-1.4018	-1.5264	-1.6431	5.8072	5.6280	3.8707
	(6.17)***	(2.38)**	(3.95)***	(1.00)	(2.35)**	(9.11)***	(3.91)***	(3.78)***	(4.54)***	(5.62)***	(5.28)***	(8.90)***
Adj. R^2	0.88	0.59	0.83	0.71	0.65	0.69	0.69	0.84	0.87	0.92	0.92	0.94
Number of obs.	2,092	2,092	2,092	2,092	2,092	2,092	2,059	2,092	2,059	2,059	2,059	2,059
Impact of market risk	0.9186	0.8011	0.9338	-1.1264	0.8443	0.7992	1.1693	0.9309	1.0231	0.9445	0.9519	0.0845
Impact of liquidity risk	-0.0086	-0.2236	-0.0160	0.0190	-0.1869	-0.1822	-0.0368	-0.0346	0.0344	0.0266	0.0200	0.1474
Impact of age	0.4171	-	0.4590	-	-	-	-	0.4271	0.4355	0.4070	0.3886	0.3149
Impact of issuance	-0.0903	-	-	-	-	-0.1247	-	-	-	-0.0899	-0.0876	-0.0627

Table 2.6
Monthly TIPS yields and illiquidity – Market integration

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of markets being integrated. This implies that the market return is the equally weighted average of assets in zero net supply, thus indexed and nominal Treasuries. The dependent variable is the TIPS yield. Market and illiquidity betas are estimated based on Equation 2.4 as loadings on the market and the non-traded illiquidity factors. We define age as the years passed since issuance, and time-to-maturity as the years until maturity. The latter captures term structure effects. The on-the-run variable is a dummy that equals one of an issue is the newest of its tenor and zero otherwise. We use the natural logarithm of the original issued amounts; whereas (lagged) yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is take over the different maturities for a given month. The economic impact is captured by the interquartile spread: we multiply the coefficient by the difference between the betas that correspond to the first and third quartile in the cross-section of betas. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions where errors take into account the averaging and include a 12-lag Newey-West correction. The sample period is July 2004 until December 2011. t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Market beta	0.7800	0.7360	0.7910	-1.0904	0.7594	0.7320	0.9124	0.8112	0.8530	0.7996	0.8335	-0.0234
	(3.87)***	(3.30)***	(4.04)***	(12.81)***	(3.44)***	(3.23)***	(4.41)***	(3.95)***	(5.46)***	(5.19)***	(5.10)***	(-0.28)
Illiquidity beta	-0.0394	-0.3557	-0.0428	-0.0462	-0.3140	-0.2842	-0.2094	-0.0356	0.0222	0.0134	0.0428	0.2034
	(-0.55)	(3.94)***	(-0.54)	(-0.30)	(3.33)***	(3.24)***	(3.18)***	(-0.45)	(0.32)	(0.21)	(0.62)	(2.48)**
Age	0.0898		0.1007					0.0953	0.0982	0.0894	0.0876	0.0690
	(8.87)***		(11.40)***					(6.27)***	(11.79)***	(9.12)***	(6.55)***	(6.39)***
Time-to-maturity				0.1694								0.0883
				(7.36)***								(3.81)***
On-the-run indicator					-0.2235			0.0576			0.0578	0.0366
					(2.56)**			(0.58)			(0.72)	(0.74)
Issued amount	-0.2930					-0.3933				-0.2957	-0.2875	-0.2268
	(8.09)***					(10.87)***				(7.62)***	(7.02)***	(9.95)***
Yield volatility$_{t-1}$							4.7524		2.0170	0.7822	1.0465	2.5372
							(3.67)***		(1.36)	(0.51)	(0.67)	(1.83)*
Intercept	5.6259	-0.6874	-1.3082	-0.5694	-0.6802	8.5267	-0.9506	-1.3234	-1.4020	5.6377	5.3974	4.0676
	(7.10)***	(-1.63)	(3.36)***	(-1.34)	(-1.56)	(9.39)***	(2.59)**	(3.45)***	(3.78)***	(6.38)***	(5.86)***	(9.10)***
Adj. R^2	0.89	0.67	0.84	0.75	0.71	0.75	0.72	0.85	0.87	0.92	0.92	0.94
Number of obs.	2,092	2,092	2,092	2,092	2,092	2,092	2,059	2,092	2,059	2,059	2,059	2,059
Impact of market risk	0.8925	0.8422	0.9051	-1.2477	0.8690	0.8376	1.0441	0.9282	0.9761	0.9151	0.9538	-0.0268
Impact of liquidity risk	-0.0264	-0.2385	-0.0287	-0.0310	-0.2106	-0.1906	-0.1404	-0.0239	0.0149	0.0090	0.0287	0.1364
Impact of age	0.4121	-	0.4623	-	-	-	-	0.4372	0.4506	0.4104	0.4020	0.3166
Impact of issuance	-0.0844	-	-	-	-	-0.1133	-	-	-	-0.0852	-0.0828	-0.0653

Table 2.7
Monthly nominal Treasury yields and liquidity – Market segmentation

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of markets being segmented. This implies that the market return is the equally weighted average of nominal Treasuries. The dependent variable is the nominal Treasury yield. Market and illiquidity betas are estimated based on Equation 2.4 as loadings on the market and the non-traded illiquidity factors. We define age as the years passed since issuance, and time-to-maturity as the years until maturity. The latter captures term structure effects. The on-the-run variable is a dummy that equals one of an issue is the newest of its tenor and zero otherwise. We use the natural logarithm of the original issued amounts; whereas (lagged) yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is take over the different maturities for a given month.The economic impact is captured by the interquartile spread: we multiply the coefficient by the difference between the betas that correspond to the first and third quartile in the cross-section of betas. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions where errors take into account the averaging and include a 12-lag Newey-West correction. The sample period is July 2004 until December 2011. t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	Benchmark	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Market beta	1.4434	1.4677	0.8223	1.4425	1.4461	1.3114	0.7446	0.7585	0.7093	0.7391	0.7280	0.7402
	(3.92)***	(3.90)***	(1.91)*	(3.92)***	(3.92)***	(3.89)***	(1.91)*	(1.99)**	(1.85)*	(2.17)**	(2.22)**	(2.25)**
Illiquidity beta	0.4913	0.4220	0.6758	0.4852	0.4763	0.3993	0.6321	0.6429	0.6532	0.5452	0.5566	0.5594
	(2.99)***	(3.30)***	(2.94)***	(3.05)***	(3.42)***	(3.76)***	(3.38)***	(3.35)***	(3.59)***	(4.27)***	(4.45)***	(4.37)***
Age		-0.0128					-0.0110	-0.0117	-0.0106	-0.0091	-0.0091	-0.0091
		(1.97)*					(1.88)*	(1.94)*	(2.12)**	(1.72)*	(2.00)**	(2.00)**
Time-to-maturity			0.0558				0.0657	0.0650	0.0686	0.0551	0.0555	0.0544
			(1.98)*				(2.79)***	(2.77)***	(2.93)***	(3.49)***	(3.58)***	(3.39)***
On-the-run indicator				0.0268				-0.0116				0.0031
				(1.16)				(-0.74)				(0.24)
Issued amount					0.0269				-0.0047		-0.0103	-0.0143
					(0.60)				(-0.25)		(-0.57)	(-0.85)
Yield volatility$_{t-1}$						3.6673				2.3864	2.5547	2.4881
						(1.74)*				(1.22)	(1.29)	(1.26)
Intercept	0.1528	0.2003	0.1779	0.1515	-0.4904	0.2535	0.2435	0.2468	0.3551	0.3385	0.5862	0.6833
	(0.70)	(0.89)	(0.96)	(0.70)	(-0.48)	(1.25)	(1.37)	(1.37)	(0.66)	(1.86)*	(1.09)	(1.31)
Adj. R^2	0.89	0.94	0.94	0.89	0.92	0.94	0.96	0.96	0.96	0.97	0.97	0.97
Number of obs.	2,132	2,132	2,132	2,132	2,132	2,100	2,132	2,132	2,132	2,100	2,100	2,100
Impact of market risk	1.8322	1.8631	1.0437	1.8311	1.8356	1.6646	0.9452	0.9628	0.9003	0.9382	0.9241	0.9396
Impact of liquidity risk	0.1313	0.1128	0.1806	0.1296	0.1273	0.1067	0.1689	0.1718	0.1745	0.1457	0.1487	0.1495

Table 2.8

Monthly nominal Treasury yields and liquidity – Market integration

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of markets being segmented. The dependent variable is the nominal Treasury yield. Market and liquidity betas are estimated based on Equation 2.5 as loadings on the market and the non-traded liquidity factors. We define age as the days passed since issuance, and time-to-maturity as the days until maturity. The latter captures term structure effects. The on-the-run variable is a dummy that equals one of an issue is the newest of its tenor and zero otherwise. We use the natural logarithm of the original issued amounts; whereas (lagged) yield volatility is defined as the difference between the standard deviations of individual issues and the cross-sectional average standard deviation of quoted yields, where the average is take over the different maturities for a given month. The economic impact is captured by the interquartile spread: we multiply the coefficient by the difference between the betas that correspond to the first and third quartile in the cross-section of betas. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions where errors take into account the averaging and include a 12-lag Newey-West correction. The sample period is July 2004 until December 2011. t-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Market beta	1.4396	1.4751	0.9771	1.4551	1.4533	1.2287	0.8515	0.8441	0.9433	0.7475	0.8331	0.8375
	(4.01)***	(4.09)***	(2.13)**	(3.97)***	(4.09)***	(4.11)***	(2.09)**	(2.06)**	(2.06)**	(2.55)**	(2.51)**	(2.50)**
Illiquidity beta	0.2414	0.1940	0.4578	0.2659	0.2559	0.1235	0.4497	0.5266	0.5028	0.3137	0.3510	0.3784
	(2.38)**	(2.44)**	(2.26)**	(2.23)**	(3.32)***	(2.48)**	(2.46)**	(2.60)**	(2.81)***	(2.58)**	(3.01)***	(3.09)***
Age		-0.0175					-0.0157	-0.0168	-0.0215	-0.0117	-0.0164	-0.0161
		(2.26)**					(2.13)**	(2.04)**	(2.53)**	(1.94)*	(2.45)**	(2.34)**
Time-to-maturity			0.0478				0.0639	0.0686	0.0577	0.0529	0.0471	0.0482
			(1.53)				(2.38)**	(2.51)**	(2.09)**	(3.27)***	(2.90)***	(2.95)***
On-the-run indicator				-0.0047				-0.0766				-0.0124
				(-0.09)				(-1.01)				(-0.37)
Issued amount					0.0180				-0.1040		-0.0799	-0.0781
					(0.35)				(3.52)***		(4.37)***	(4.40)***
Yield volatility$_{t-1}$						5.4332				4.7186	4.6556	4.3847
						(2.03)**				(1.87)*	(1.87)*	(1.87)*
Intercept	0.2270	0.2771	0.2119	0.2202	-0.2042	0.3685	0.2909	0.2929	2.7727	0.4521	2.3547	2.3047
	(1.09)	(1.26)	(1.30)	(1.06)	(-0.19)	(1.90)*	(1.77)*	(1.77)*	(4.26)***	(2.78)***	(5.81)***	(6.07)***
Adj. R^2	0.88	0.93	0.92	0.88	0.91	0.93	0.95	0.95	0.95	0.97	0.97	0.97
Number of obs.	2,132	2,132	2,132	2,132	2,132	2,100	2,132	2,132	2,132	2,100	2,100	2,100
Impact of market risk	1.8494	1.8950	1.2553	1.8694	1.8671	1.5786	1.0939	1.0845	1.2119	0.9603	1.0703	1.0759
Impact of liquidity risk	0.1285	0.1032	0.2436	0.1415	0.1362	0.0657	0.2393	0.2803	0.2676	0.1669	0.1868	0.2014

Table 2.9
Liquidity regimes

The table reports estimates for the second step (Equation 2.5) of the Fama-MacBeth procedure under the assumption of markets being segmented, but considering months of increasing and decreasing liquidity separately. The dependent variable is either the bond yields or the swap. Market and liquidity betas are estimated based on Equation 2.4 as loadings on the market and the non-traded liquidity factors. We define age as the days passed since issuance, and time-to-maturity as the days until maturity. To proxy for size, we use the natural logarithm of the original issued amounts. Displayed coefficients are average figures from the monthly repeated cross-sectional regressions where errors take into account the averaging. The sample period is July 2004 until December 2011. T-statistics are given in parentheses and *, **, and *** denote significance at the 10%, 5%, and 1% level.

Panel A: Monthly TIPS yields

	Benchmark	Increasing illiquidity	Decreasing illiquidity
Segmented market beta	0.932	0.892	0.975
	(10.08)***	(6.33)***	(8.22)***
Segmented illiq. beta	-0.032	-0.048	-0.013
	(0.88)	(0.85)	(0.31)
Age	0.088	0.090	0.085
	(23.07)***	(14.70)***	(19.71)***
Size	-0.239	-0.247	-0.230
	(19.35)***	(13.91)***	(13.39)***
Constant	4.241	4.605	3.844
	(13.01)***	(9.88)***	(8.50)***
R^2	0.87	0.86	0.88
N	2,216	1,155	1,061

Panel B: Monthly nominal Treasury yields

	Benchmark	Increasing illiquidity	Decreasing illiquidity
Segmented market beta	1.372	1.421	1.311
	(13.41)***	(10.17)***	(8.66)***
Segmented illiq. beta	0.524	0.537	0.508
	(8.05)***	(5.96)***	(5.35)***
Constant	0.230	0.244	0.214
	(3.49)***	(2.56)**	(2.37)**
R^2	0.89	0.89	0.89
N	2,252	1,255	997

Panel C: Monthly inflation swap rates

	Benchmark	Increasing illiquidity	Decreasing illiquidity
Segmented market beta	0.534	0.555	0.511
	(10.46)***	(7.12)***	(7.76)***
Segmented illiq. beta	0.017	0.018	0.017
	(6.41)***	(4.37)***	(4.71)***
Constant	1.940	1.917	1.965
	(18.54)***	(11.79)***	(14.88)***
R^2	0.72	0.71	0.72
N	1,350	690	660

Table 2.10
The mispricing

This table presents descriptive statistics of the replicated trading strategy of Fleckenstein et al. (2014) (FLL). In this strategy we compare the prices of a nominal Treasury issue to its replicating portfolio that consist of a maturity matched TIPS issue, inflation swap contracts and STRIPS issues. Panel A presents the results of the replication for our sample, alongside with the two cases of liquidity corrections applied to this strategy. The correction is based on adjusting yields with estimated liquidity premiums from Equation 2.5, both under the assumption of the three markets being segmented and integrated. Panel B exhibits the difference between the original strategy and the adjusted versions, where the difference is defined as the FLL mispricing minus the corrected series. The data correspond to 26 bond pairs in the sample period between July 2004 and December 2011.

Panel A: Mispricing and corrections

	Mean	St. Dev.	Min	p25	p75	Max
FLL mispricing	3.07	1.69	0.42	2.22	3.32	11.80
Segmentation-corrected price differential	0.77	1.62	-1.64	-0.07	0.95	9.20
Integration-corrected price differential	0.39	1.78	-2.12	-0.52	0.68	9.64

Panel B: The effect of liquidity correction

	Mean	St. Dev.	Min	p25	p75	Max
Difference in segmentation	2.31	0.21	1.80	2.19	2.45	2.90
Difference in integration	2.68	0.31	1.97	2.47	2.83	3.79

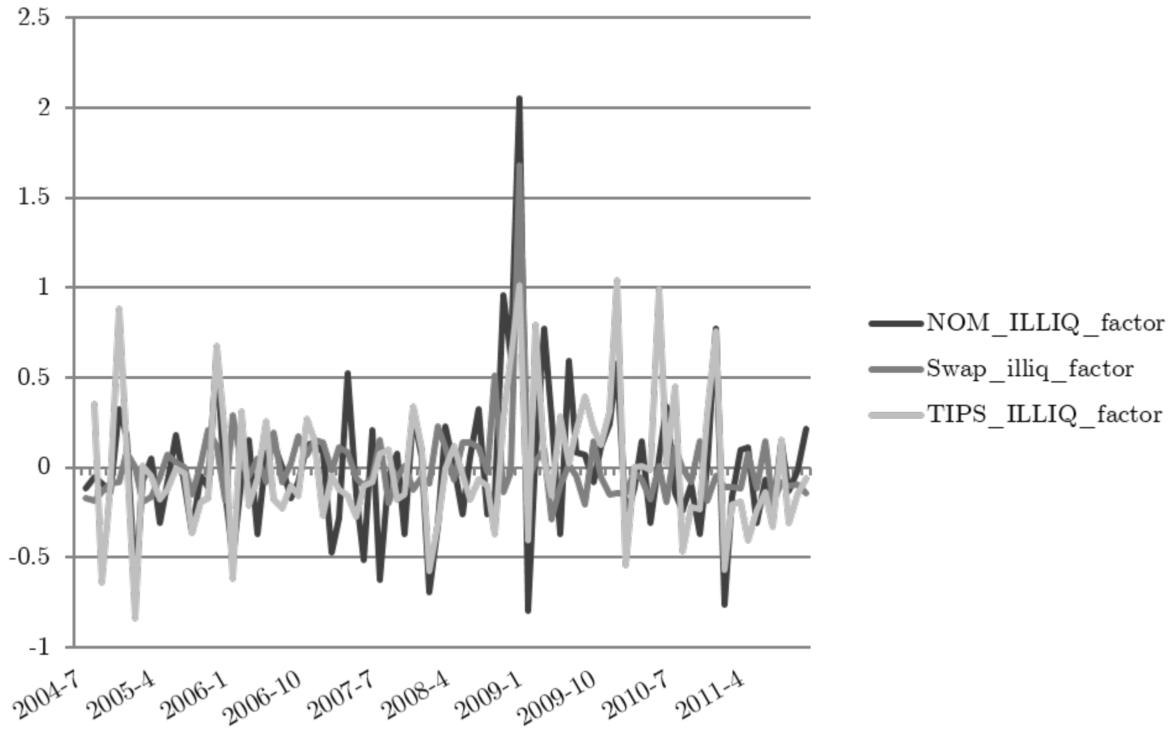
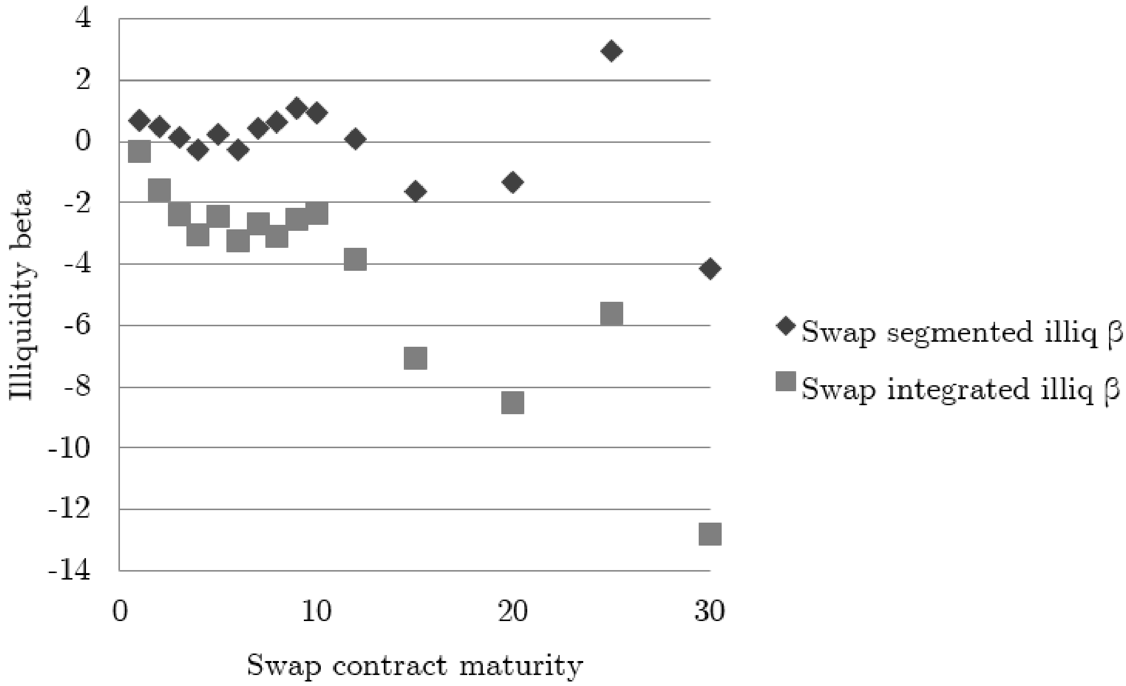


Figure 2.1 Illiquidity factors

The figure depicts the time evolution of the non-traded illiquidity factors. They are residuals from autoregressive processes: AR(3) for ILLIQ measures of nominal and indexed bonds, and AR(1) for the average Roll measure of inflation swaps.



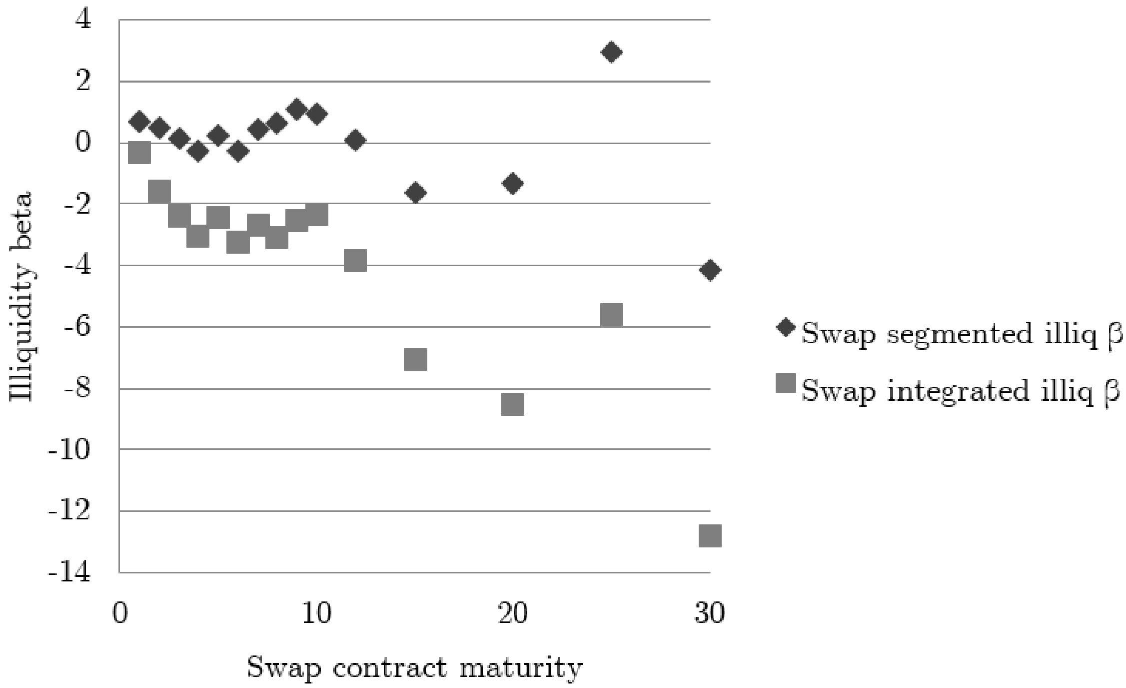


Figure 2.2 Inflation swap market and illiquidity betas

The scatter plots depict the betas estimated from the time series regressions of inflation swap returns on market and illiquidity factors. The above plot focuses on illiquidity betas, whereas the lower panel presents market betas, both estimated under the assumption of market segmentation and integration.

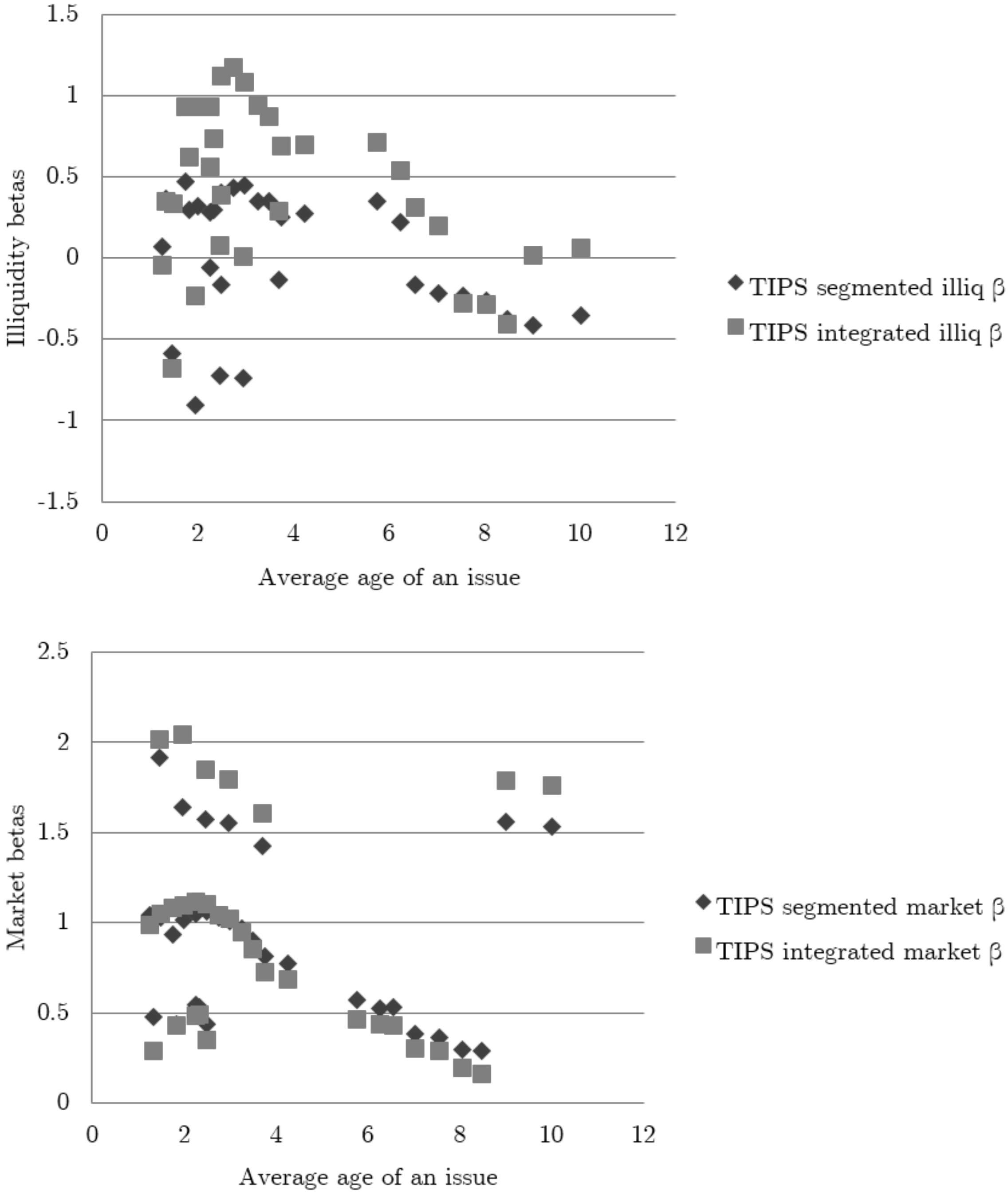
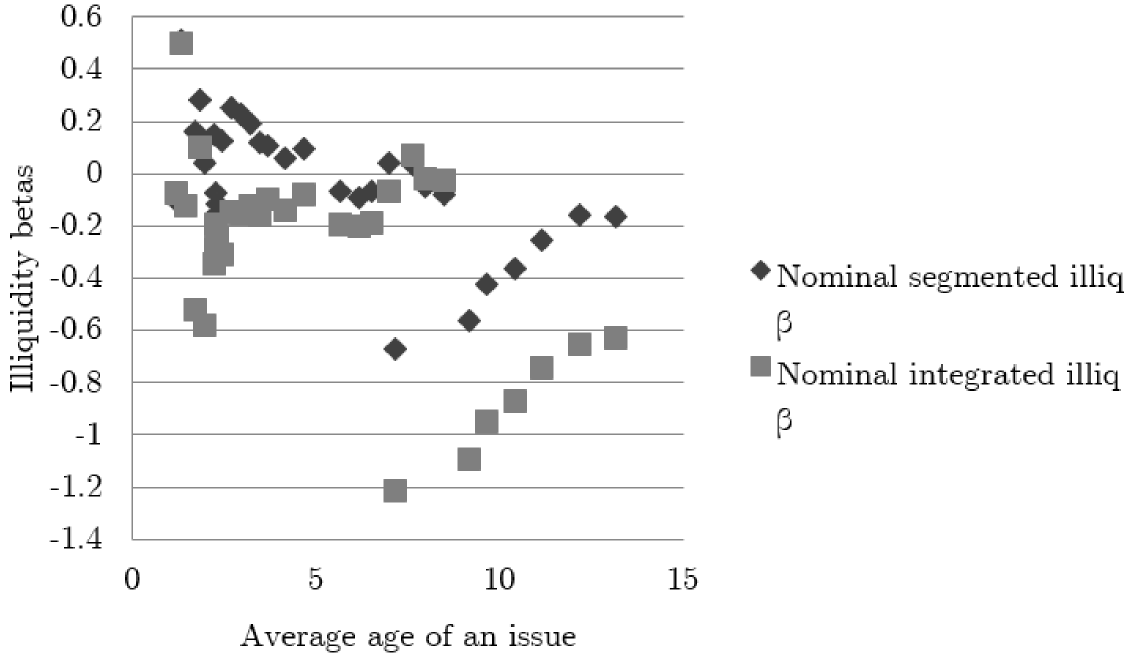


Figure 2.3 TIPS market and illiquidity betas

The scatter plots depict the betas estimated from the time series regressions of TIPS returns on market and illiquidity factors. The above plot focuses on illiquidity betas, whereas the lower panel presents market betas, both estimated under the assumption of market segmentation and integration. Betas are sorted on average age of bond issues.



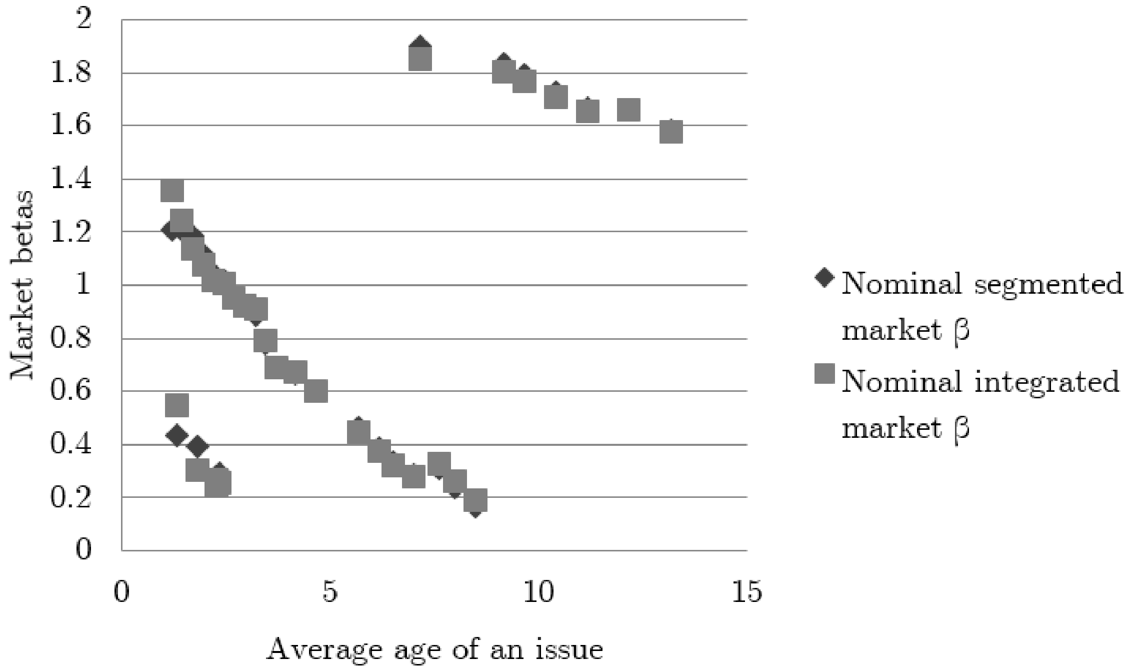


Figure 2.4 Nominal Treasury market and illiquidity betas

The scatter plots depict the betas estimated from the time series regressions of nominal Treasury returns on market and illiquidity factors. The above plot focuses on illiquidity betas, whereas the lower panel presents market betas, both estimated under the assumption of market segmentation and integration. Betas are sorted on average age of bond issues.

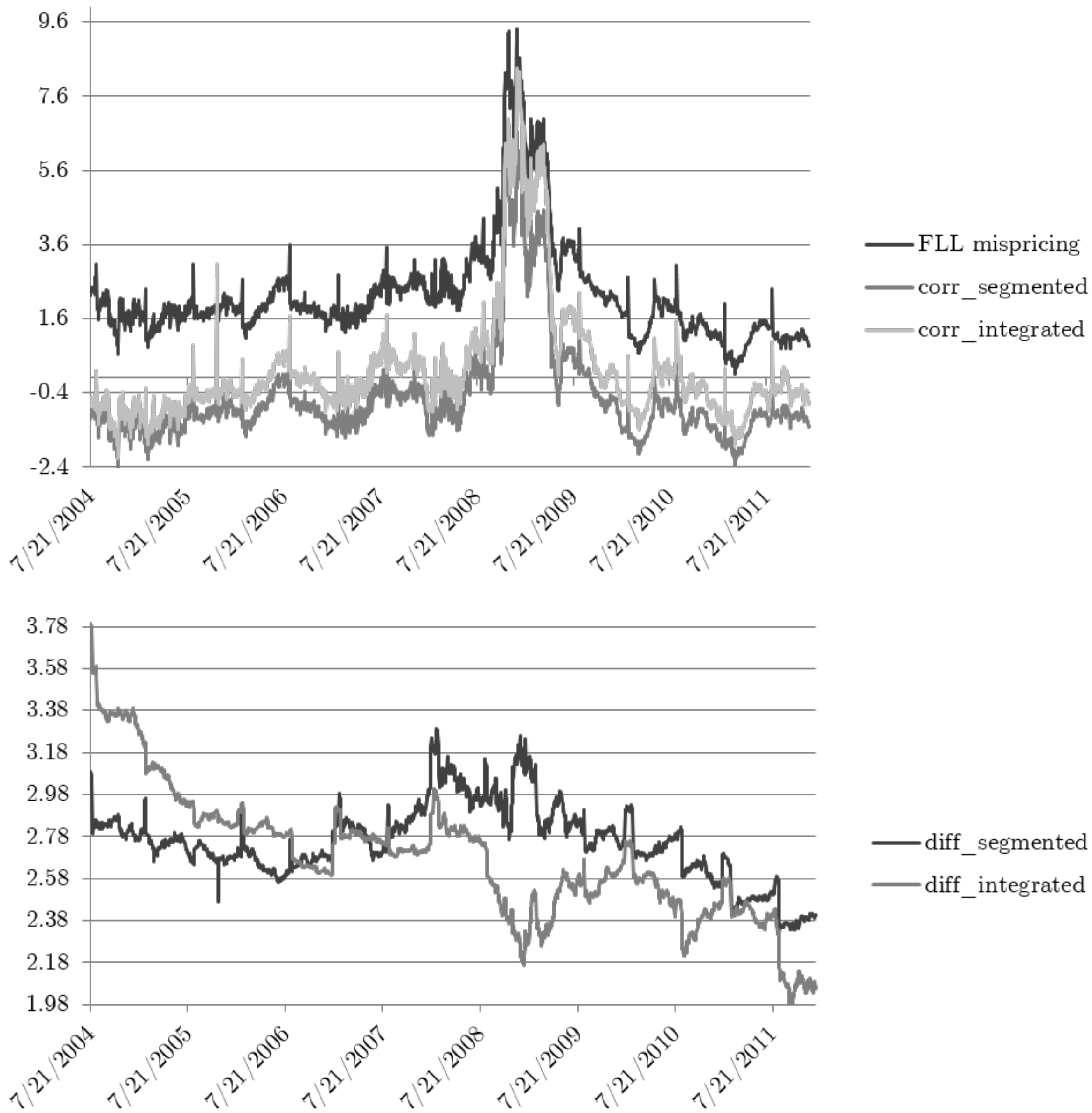


Figure 2.5 Mispricing and the effect of liquidity correction

The figure depicts the time-series behavior of the equally weighted average mispricing series across 26 maturity matched bond pairs, based on Fleckenstein et al. (2014). The mispricing is the price difference between a nominal Treasury issue and its replicating portfolio that consists of a maturity matched TIPS issue, inflation swap contracts and STRIPS issues. This panel also depicts the liquidity-adjusted series, where corrections based on segmented or integrated markets are applied. In the lower panel, we show the difference between the replicated and liquidity-corrected series.

Chapter 3

Much ado about nothing: A study of differential pricing and liquidity of short and long term bonds

3.1 Introduction

European pension funds and insurers managed more than €3.5 trillion worth of assets in 2015. For these institutions it is crucially important to attain precise estimates of long term discount rates for their asset management and valuation of liabilities for regulatory purposes. Despite its practical importance and potential welfare consequences, modelling and examining the long end of the nominal term structure has attracted little attention in the academic literature.

This paper aims to fill this gap by studying the differential pricing of short and long maturity bonds, especially focusing on segmentation in yields and liquidity. To address this question, we explore the channels through which this affects the pricing of short and long ends of the German nominal term structure between 2005 and 2015. In this period, it is conceivable that there was investor segmentation or formation of clienteles due to regulation-induced demand pressure or the effect of unconventional monetary policy on yields. Liquidity differences might have arisen during the financial crisis, while during the euro crisis potential credit premium differences could have emerged, alongside with the effect of safe haven flows of Eurozone investors. The confounding presence of the fore

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mentioned forces makes this market and time period an ideal testing ground to study whether these events have a differential effect on long and short maturity bonds.

The key result of our paper is that although there are statistically significant differences in the pricing and drivers of short and long maturity bonds, the corresponding economic effects are rather small. This means that long yields are not extensively distorted by demand pressure, default or liquidity premiums, therefore there is little evidence for substantial yield segmentation. This finding has important policy implications. Part of the policy discussion on valuation of pension and insurance liabilities is how to model long term discount rates. The current approach is based on the ultimate forward rate, which is an extrapolation method used to calculate discount rates for maturities beyond twenty years. For longer maturities, there is a variety of methods that use interest rate swaps. However, in light of our results, this practice seems unnecessary: if long maturity bond yields are not distorted, we could extrapolate long term discount rates from these yields observed in bonds markets. Is the ultimate forward rate discussion much ado about nothing? One the one hand, there is a lot of money at stake, thus making sure that this wealth is properly valuated is crucial. On the other hand, it seems that we can trust yields of long maturity bonds, as any existing effects of yield segmentation are negligibly small.

Our second key finding is that we present evidence for some degree of liquidity segmentation across short and long maturities. This finding is in line with the theory of liquidity segmentation (Amihud and Mendelson, 1986; Beber et al., 2012), however due to the small cross-section of long maturity bonds, we cannot formally test price segmentation in long and short yields. Nevertheless, we show that the nature of liquidity varies along the curve: liquidity of short maturity bonds seems more systematic in nature, whilst liquidity of long maturity bonds behaves independently from other market measures. These results seem plausible as long maturity bonds are likely to be held by long horizon buy-and-hold investors, who might be inherently less concerned about the issue level illiquidity and cost of trading, as opposed to investors of shorter maturity bonds.

Our empirical approach is as follows. First, to study segmentation, we construct a pair of measures similar to that of Hu et al. (2013). The noise measure is a liquidity proxy that is based on the pricing errors of observable yields compared to a smooth theoretical yield curve. We fit a Nelson-Siegel model to bonds up to twenty years of maturity to determine the theoretical curve, which we extrapolate for longer maturities.[1] Next, we construct two measures of noise that allow us to focus on deviations of long and shorter

[1] By using this method, we cannot exclude the possibility that the whole curve could be affected by some bias. However, we claim and provide empirical evidence that the long end of the yield curve is relatively more distorted.

maturities separately: the average root mean squared pricing error for short maturities is 7.91 basis points, while it is 19.39 basis points for maturities longer than 20 years. The size of long noise measure suggests that with even such a simple model we get good predictions of long maturity bond yields. We find that the two measures capture different aspects of liquidity: short noise is related to the age of the bond and to market and funding liquidity; while long noise is linked to bond issue level illiquidity and the US noise measure. Nevertheless, the low explanatory power of conventional liquidity proxies suggests that the two noise measures capture an aspect of illiquidity that correlates, but goes beyond traditional liquidity measures.

Next, to further examine the pricing of bonds, we also look at the average fitting error of long maturity bonds, called the bias. This fitting error is on average -9.25 basis points and is persistently negative, meaning that the observed long yields are consistently below those implied by our pricing model, although this difference economically quite small especially given the simple model we fit for the yield curve. In search for the reason why long yields are too low', we find that the bias is linked to the Roll implied bid-ask spread, time to maturity of bonds, liquidity risk and the Ted spread, flight-to-safety flows and the credit risk of Germany. To deepen our understanding of the time-series drivers of short and long maturity bond yields, we study their exposures to proxies of demand pressure, default premium and liquidity. The yield decomposition reveals little evidence of significant segmentation in German sovereign bond markets. Short yields are driven by short noise and large scale asset purchases; the latter is indirect evidence that these unconventional monetary measures have an effect on yields of targeted maturities. Additionally, we find that both long and short yields are strongly linked to German credit quality and flight-to-safety flows. This suggests that when periphery credit quality declines, the overall German yield curve becomes a target to safe haven flows of euro zone investors. Although these results are statistically significant, the economic effects, measured by one standard deviation change of the explanatory variables, are in the magnitude of 1-3 basis points.

Looking at liquidity segmentation, we study how yields of the short and long ends of the yield curve are linked to a wide range of liquidity measures. The effect of short noise on the average short yield is significant and robust to the inclusion of other measures, moreover, it varies over time. As opposed to this, we find that the average long yield has a weaker connection to long noise, while we find evidence of liquidity spillover from the short to the long end of the yield curve. We also consider if short and long noise are related to different market forces. Short noise is sensitive to large scale asset purchases of the ECB, as well as to changes in stock market volatility and flight-to-safety flows. The long noise measure has a similar relationship to safe haven flows. However, the fitting errors of long maturity bonds are also sensitive to declining credit quality of the issuer, as well as to breakup risk.

The remainder of the chapter is organized as follows. Section 3.2 discusses the related literature, while Section 3.3 explains the channels of yield and liquidity segmentation in bond markets. Section 3.4 presents the data, and the German nominal sovereign bond market and elaborates on how we measure the effect of bond market segmentation by the two noise measures. After establishing noise as a proxy for liquidity, Section 3.5 introduces the bias and presents the results of the decomposition of short and long yields, as well as the drivers of yield and liquidity segmentation. Finally, Section 3.6 discusses policy implications of our results and Section 3.7 concludes.

3.2 Literature overview

This paper studies the differential pricing of short and long maturity bonds, especially focusing on segmentation in yields and liquidity. But why would there be differences that are maturity dependent? In answering this question, we build on previous research on market segmentation, limits of arbitrage, liquidity and credit risk of sovereign bonds, the term structure of liquidity premium, flights-to-safety and the effects of unconventional monetary policy on sovereign bond yields. Nonetheless, the distinctive feature of our analysis is that we take a comprehensive approach of examining yields and liquidity features of bonds from a wide range of the maturity spectrum. In particular, we focus on bonds with tenors between 2 to 30 years to study the differential effect and channels of demand pressure, default risk and liquidity along the yield curve. To do so, we use a method closely related to Hu et al. (2013).

Hu et al. (2013), henceforth HPW, develop a measure of bond liquidity that is based on the pricing error of observable yields compared to a theoretical smooth curve. They show that this measure is linked to the amount of arbitrage capital on the market: the pricing error is larger and more persistent when funding liquidity conditions deteriorate and arbitrageurs lack the resources to trade substantial price deviations away. We construct a similar measure, but instead of aggregating all information along the yield curve, we focus on deviations of long and shorter maturities separately. These two measures of noise help us explore how prices and liquidity can differ across maturities. First, some investors might not have access to the overall yield curve due to either regulatory constraints or endogenous choice regarding their investment horizon. This gives rise to market segmentation. Second, if local demand shocks are accompanied by limited arbitrage capital, pricing inefficiencies cannot be eliminated and become persistent. Third, if short and long maturity bonds are exposed to risk factors to a different extent, different levels of risk premiums will emerge.

Addressing the first explanation, we contribute to the literature on market segmentation. Among other factors, such as time-varying expectations of future rates, changes in bonds yields can be attributed to changes in demand or supply. On the one hand, bond demand might be driven by preferred habitat investors. Vayanos and Vila (2009) show that if investors have preferences for specific maturities, bond markets become segmented, in which case demand shocks affect the cross-section of bonds differently. Greenwood and Vayanos (2010) provide empirical evidence for a preferred habitat induced demand shock following the 2004 UK pension reform. Related to this, our study also aims to deepen the understanding of how demand pressure due to regulatory changes, specifically those concerning the natural clientele of the bonds in our sample, affect yields and liquidity in the cross-section of maturities. On the other hand, supply factors also play a role in the determination of yields.Greenwood and Vayanos (2014) and Guibaud et al. (2013) explain and present models of how supply and maturity structure of sovereign debt influences bond yields.

Another channel through which local demand pressure arises is unconventional monetary policy.[2] While Krishnamurthy and Vissing-Jorgensen (2011) show that quantitative easing (QE) in the US caused a supply shortage for certain clientele demand, pushing yields downwards, Christensen and Gillan (2013) provide evidence on how market liquidity is improved by such measures. D'Amico et al. (2012) present similar evidence on the effect of large-scale asset purchase programs (LSAP) on the preferred habitat and duration of sovereign bonds, alongside with D'Amico and King (2013) who show that LSAP often causes market segmentation. There is also growing evidence from the intervention of the European Central Bank (Eser and Schwaab, 2016; Krishnamurthy et al., 2015) and its effect on the liquidity of European sovereign bond markets (De Pooter et al., 2013; Pelizzon et al., 2014, minga). Our contribution to this literature lies in looking at how yields and liquidity of bonds with different maturities are affected by these measures and how this could lead to structural distortions of the yield curve.

We also consider the second reason for differences. In a well-functioning market, distortions local to certain maturities could be traded away by arbitrageurs constantly searching for such mispricing to profit from. However, market frictions, for instance margins, haircuts and other constraints, might prevent such investors from trading on these inefficiencies, as the limits of arbitrage literature describes (Shleifer and Vishny, 1997; Gromb

[2]A study that does not specifically focus on unconventional measures, but buyback programs is that of Phillips (2003), who documents significant bond market reactions to Treasury announcements of reduced supplies of 30-year bonds, only local to the yields of those bonds.

and Vayanos, 2002; Liu et al., 2006; Gorton and Metrick, 2010; Ashcraft et al., 2010). Additionally, arbitrage capital might move too slowly to eliminate such opportunities, due to capital constraints or agency problems of delegated asset management, as in Mitchell et al. (2007); Duffie (2010).

The third driver leading to differential pricing could be alternate levels of risk premiums in short and long term yields. We consider the following premiums in our analysis: flight-to-safety premium, liquidity premium, credit risk premium and partial default or breakup risk premium. By studying whether short and long maturity bonds have different sensitivity to flight-to-safety flows, we also enrich the literature on this component of bond demand. These episodes occur when due to a credit shock, investors reallocate their portfolios towards safer assets. This can happen both within, as well as across asset classes (Næs et al., 2011) and is typically accompanied by flights-to-liquidity (Longstaff, 2004). Beber et al. (2009) disentangle flight-to-safety and liquidity episodes using a sample of Eurozone bonds from 10 countries, whereas Baele et al. (2015) develop a methodology to identify such episodes in a sample of 23 countries. Ejsing et al. (2012) identify safe haven flows in German and French bond markets. We also provide evidence that when periphery credit quality declines in the Eurozone, the entire German yield curve, irrespective of maturities, becomes a target for such flows.

Further, our paper contributes to the existing literature on liquidity and the term structure of liquidity premium in bonds. Liquidity premium of government bonds can be analyzed along two dimensions. Early studies typically estimate liquidity premium by identifying pairs of assets that are otherwise similar, but have different levels of liquidity. These assets can either have the same issuer and differ in their seasonedness (Krishnamurthy, 2002; Gürkaynak et al., 2007; Bühler and Vonhoff, 2011; Schuster and Uhrig-Homburg, 2013; Kempf et al., 2012; Fontaine and Garcia, 2012) or provide the same guarantees while being issued by different institutions (Longstaff, 2004; Krishnamurthy and Vissing-Jorgensen, 2011; Schuster and Uhrig-Homburg, 2013; Schwarz, 2015). However, this latter measure is often criticized as it not only captures liquidity premium, but also breakup or partial default risk.

Much less is known about liquidity differences in the cross-section of different bond tenors, especially of the very long maturities. Most papers that study this question focus on the US corporate bond market: Ericsson and Renault (2006) find a decreasing term structure, while Dick-Nielsen et al. (2012) find that liquidity premium increases with maturity. Gehde-Trapp et al. (2016) provide a third view: they find a U-shaped premium, where the very short and very long maturities carry the largest compensation for illiquidity. Goyenko et al. (2011) study liquidity premium differences and explain the on- and off-the-run premiums of three maturity buckets: short maturity T-bills, 2-5 year and 10-year bonds. Despite its practical importance, the long end of the term structure has attracted much less attention in academic research. Our study aims to fill this gap, by showing how yields and liquidity can substantially differ among longer and shorter maturities of nominal German sovereign bonds.

And finally, this paper also relates to the strand of literature that separates liquidity and credit effect in sovereign yields. Beber et al. (2009) disentangle the effects of liquidity and credit quality in 10 Eurozone countries. Similarly, Schwarz (2015) separates the components of yields due to liquidity and credit risk. Ejsing et al. (2012) quantify liquidity and credit risk premiums in German and French government bond yields based on a state-space model with two latent factors. Bai et al. (2012) examine what caused the recent sovereign bond crisis illiquidity of markets or deteriorating credit conditions. Darbha and Dufour (2014) study the term structure of default and illiquidity in a sample of nominal Euro area government bonds, whereas Monfort and Renne (2014) present an arbitrage-free model with joint dynamics of euro-area bond spreads that are driven by liquidity and credit risk. Besides, we not only consider the sovereign risk of the issuer, but test for breakup or partial default risk component in German yields, as in De Santis (2015); Simon (2015).

3.3 Bond market segmentation

The expectations hypothesis of the term structure (EHTS) postulates that risk premiums on long-term bonds, the expected excess returns of long maturity bonds over their short ones, should be constant over time. If EHTS holds, then risk premiums would not vary over time. Consequently, as Pflueger and Viceira (2011) point out, investors could not earn predictable returns on switching between long and short maturities. As such, bonds of short and long maturities should be close to perfect substitutes and the shape of the yield curve should only depend on investors' expectation of future rates.

However, there is ample evidence of predictable and time-varying risk premiums in bonds (Bliss and Fama, 1987; Cochrane and Piazzesi, 2005; Pflueger and Viceira, 2015). Moreover, in reality different maturities are not perfect substitutes due to the presence of investor clienteles with preferences for specific maturities. As a consequence, the interest rate for a given maturity is also influenced by demand and supply shocks specific to that maturity. Vayanos and Vila (2009) show that if investors have preferences for certain maturities, bond markets become segmented. Additional evidence for the presence of preferred habitat investors is presented by Greenwood and Vayanos (2010), who find that the average maturity of government debt predicts excess bond returns. On the other hand, Guibaud et al. (2013) show that catering to maturity clienteles is an optimal issuance policy, as a welfare-maximizing government issues longer maturity debt when the fraction of long to short horizon investors increases. Additional evidence from the corporate bond market also suggests that corporations engage in gap-filling behavior, issuing long-term debt at times when the supply of long-term government debt decreases (Greenwood et al., 2015).

In light of the above evidence, we think that segmentation is a result of either regulatory constraints or endogenous choice due to which investors choose to hold only a part of the yield curve. Prominent examples of such regulation are the Basel and Solvency regulatory frameworks that incentivize banks and insurers and pension funds, respectively, to tilt their asset portfolios towards certain maturities. On the other hand, clienteles can also arise due to the costs of trading or investors' holding period. Amihud and Mendelson (1986) show that investors with different holding periods trade assets with different relative spreads, whereas the recent work of Beber et al. (2012) show that in the presence of investors with heterogeneous investment horizons, the pricing of liquidity also becomes segmented. In segmented market, short and long maturity bonds are not perfect substitutes and therefore investors cannot achieve the same risk exposures by holding only certain bond tenors. This would be otherwise in integrated markets, where all parts of the market are spanned by the same risk factors. This paper aims to discover the extent to which German sovereign bonds are subject to segmentation and the channels through which this segmentation affects the pricing of short and long maturity bonds. More specifically, we are interested in whether segmentation implies pricing differences along the yield curve, and what underlying drivers can cause these differences. We also examine how liquidity of long and short bonds differs and whether they it is priced differently. The next subsections will present the implications of segmentation on sovereign yields and on their liquidity features.

3.3.1 Segmentation in observable yields

In integrated markets, where risk premiums are the same along the yield curve, a term structure model fitted on observable yields would have an equally good fit across different maturities. However, if markets are segmented, we are likely to observe larger fitting errors on the part of the curve where yields are most distorted by segmentation. The fore mentioned fitting errors are a good measure for these structural distortions and are used in this study to proxy for segmentation in long maturity bonds. We define the bias as the average fitting error of long maturity bond yields compared to the Nelson-Siegel curve that we estimate from yields of German sovereign capital market securities.[3] The idea of using the quality of fit to proxy for segmentation effects is similar to that of HPW. They show that the mean squared fitting error of yields is a proxy for market-wide liquidity, capturing the aspect of liquidity that is closely linked to the available arbitrage capital and funding liquidity dry ups around crises.

We analyze whether the size and sign of this structural distortion in yields captured by the bias varies over time, depending on the forces that drive the short and long ends of the yields curve. To measure the differential effect of these drivers, we separately decompose short and long yields in our sample. Our yield decomposition is inspired by Krishnamurthy et al. (2015), henceforth KNVJ, who show that yields of Eurozone sovereign bonds contain the following components: an expectations hypothesis term and term premium, default risk premium, redenomination risk premium and a residual term that arises due to illiquidity frictions. While focusing on differences between short and long yields, we expand the list of potential yield components. We take a broader view on segmentation and liquidity frictions to concentrate on the following channels through which it can affect yields: demand pressure, default premium and differential liquidity of short and long bonds.

Demand pressure can arise due to regulatory changes, flight-to-safety flows and as a product of unconventional monetary policies, such as LSAPs and QE. First, regulatory changes matter if they affect only part of the investor clientele. For instance, the Basel II capital regulation of banking sector has an asset only view, and thus incentivizes banks to hold 10-year government bonds in their investment portfolios. This is because long maturity bonds are seen to be more risky and therefore they require higher capital buffers. As opposed to this, Solvency II focuses on asset and liability management of pension funds and insurers. Since these institutions have long term liabilities, which have to be hedged by minimizing the duration gap between assets and liabilities, the regulation encourages them to hold bonds with maturities longer than 20 years. To (indirectly) capture the effect of regulatory demand pressure, we include a time trend in our analysis. The second channel of demand pressure arises during the financial and euro crisis, when investors are flying to safe haven countries from the riskier periphery bonds. We capture this effect by a variable linked to changes in the default risk of periphery sovereign issuers in the Eurozone. And at last, the unconventional monetary policy measures applied during the fore mentioned crises have an effect on bonds that are part of the policies of the European Central bank (ECB), such as the Securities Markets Programme, the Outright Monetary Transactions, and the Long-Term Refinancing Operations. To capture the sum of these monetary policy actions, we control for the growth of the asset side of the ECB's balance

[3]The detailed description of how this measure is constructed can be found in the section on measuring the effect of market segmentation.

sheet.

Similar to KNVJ, we also differentiate between two types of default risk premium. Segmentation could also result in a differential effect of sovereign risk on the short and long ends of the yield curve. We proxy issuer credit risk by the 5-year USD-denominated German CDS spread. Despite that Klinger and Lando (2015) show that the CDS spread of a safe haven country, like Germany, is a poor proxy for its credit risk due to regulatory hedging requirements of investment banks, this is consensually the best proxy for sovereign risk. Nevertheless, we also consider breakup and selective default risks. We define selective default, as the government strategically defaulting on certain type of obligations, affecting only certain clientele groups. For instance, in case of insolvency the government could default exclusively on short maturity bonds, as they are due in the near future, while keeping promises on long maturity bonds mostly held by institutional investors. Selectively defaulting on bonds could also result in the exit from the Eurozone, a scenario in which the sovereign issuer would fulfill obligations in a currency other than in euros. This constitutes a risk for investors, who as a consequence would be paid in a depreciated currency. This risk is called the redenomination risk, closely linked with selective default and a direct consequence of the breakup of the Eurozone. We capture selective default and breakup risk by the KfW spread from Schwarz (2015). This spread is similar to the Refcorp spread of Longstaff (2004), and is defined as the yield difference between the 10-year KfW agency bond and the German Bund of matching maturity. Simon (2015) shows that if investors expect selective default to happen, a compensation for bearing this risk will be reflected in bond yields.

3.3.2 Segmentation in bond liquidity

Vayanos and Vila (2009) show that if investors have preferences for certain maturities, bond markets become segmented. Investors can choose to hold only a part of the yield curve due to either regulatory constraints or endogenously. The latter typically arises due to the costs of trading or investors holding period. Amihud and Mendelson (1986) show that investors with different holding periods trade assets with different relative spreads.

The concept of liquidity segmentation is motivated by investors with different holding periods, but has an effect that goes beyond relative trading costs. Beber et al. (2012) show that if investors have different investment horizons, liquidity also becomes segmented. They also show this has an effect on the pricing of liquidity, as accounting for investment horizon gives rise to cross-sectional differences in liquidity risk premiums. Moreover, it is also plausible that for investors with different horizons, different aspects of liquidity would matter: this is the reason why we find that our short and long noise measures are primarily driven by different aspects of asset, market and market liquidity risk. We provide further evidence for liquidity segmentation by constructing separate liquidity measures for the long and short ends of the curve, inspired by the noise measure of Hu et al. (2013). Our bond sample does not allow for formal cross-sectional pricing tests of liquidity, however we can investigate the exposures of yields of short and long maturity bonds to the segmented noise measures in the time series context. Moreover, we also allow for liquidity spillovers between the two ends of the yield curve.

As part of our analysis, we link the segmented noise measures to other proxies of liquidity, which can shed light on the differential nature of liquidity along the yield curve. It is conceivable that due to investment horizons, the aspect of liquidity that we capture will differ across shorter and long maturities. On the one hand, an investor with longer holding periods is not concerned about transaction costs per se, since they do not have to trade frequently. Taking this to the extreme, a buy and hold investor should not be concerned about liquidity risk either, as unless she has to liquidate, she will hold that asset until maturity. On the other hand, actively managed portfolios have to be rebalanced regularly. These investors will have a preference for more liquid and cheaper to trade assets, moreover, they will be more exposed to dry ups of market and funding liquidity. We also link the segmented measures of liquidity to other measures of market distress, volatility and default risk and presume that short and long end liquidity will correlate with these proxies to a different extent.

3.3.3 Why German sovereign bonds?

To discover the effects of market segmentation on yields and liquidity, we focus on German government securities between 2005 and 2015. This period is characterized by the financial crisis, a major liquidity event; and the euro crisis, a time period when Germany became a safe haven sovereign in the euro area. Moreover, a harmonized regulatory reform affecting the natural clientele of this securities also coincides with this period, namely the preparatory phase to the introduction of Solvency II.[4] Additionally, German bonds have been subject to large scale purchase programs improving market liquidity, and since the beginning of 2015 also to the quantitative easing (henceforth QE) of the European Central Bank. Thus in the past ten years we are likely to observe investor segmentation (preferred habitat or clientele), regulation-induced demand pressure, the

[4]The Solvency II Directive codifies and harmonizes regulation of insurers and pension funds in the European Union. Its main concern is similar to those of the Basel regulations, it regulates the amount of capital insurers and pension funds are obliged to have as reserves to decrease the risk of insolvency. Solvency II also has risk management and governance considerations, along with consumer protection, transparency and disclosure requirements.

effect of LSAP and QE on yields, alongside with flight-to-safety episodes. These events might have a differential effect on long and short maturity bonds, and the confounding presence of them makes the German bond market an ideal testing ground to study these effects of bond market segmentation.

3.4 Measuring the effect of market segmentation

3.4.1 The data

The data of this study are from multiple sources. Daily closing mid quotes of German capital market securities, bond characteristics, CDS quotes, and quotes of EONIA and Euro swap contracts of different maturities are from Bloomberg. Size of the asset side of ECB's balance sheet also comes from the same source. This information is complemented with KfW[5] agency bond quotes from Datastream, Ted spread and VIX from the FRED database of the St. Louis Fed and the value of the DAX index from the Deutsche Bundesbank. We also use data provided by the Deutsche Finanzagentur[6] on aggregate primary dealer transaction volumes.

The starting point to measure the effect of market segmentation is the noise measure, which we construct from daily cross-sections of bond returns from 2005 through 2015.[7] We focus on showing whether there is segmentation in yields, as well as in liquidity along the yield curve; therefore, we use all German sovereign nominal capital market securities in our analysis. Bonds in our sample have fixed maturities and coupons, the same market conventions and microstructure; and can be assigned to one of the three groups: Schtzes, which are 2 year notes; Bobls are 5 year notes; and Bunds are either 10 or 30 years of maturity. Similar to other studies in the yield curve literature, we drop bonds with less than 6 months to maturity due to liquidity problems, however, we do not impose an upper bound on time to maturity. This is because we are specifically interested in how yields of long maturity bonds differ from their shorter counterparts and whether different market forces drive them.

In the resulting sample we have on average a cross-section of 52 bonds, out of which

[5]KfW or Kreditanstalt fr Wiederaufbau is a German government-owned development bank in Frankfurt. Its majority owner if the Federal Republic of Germany, which explicitly guarantees bonds KfW issues in capital markets.

[6]German Finance Agency, the German equivalent of the Treasury.

[7]Our dataset starts at January 2000 but we choose to restrict our analysis to the last ten years of the sample, where not only data quality is better, but controls for identifying the effect of market segmentation on yields are available.

on average 5 bonds are of maturities longer than 20 years. However, the cross-section varies over time: its size ranges between 41 and 61 issues of capital market securities, with 1 to 8 long maturity bonds at a time. The average maturity of bonds is 8.34 years, whereas their average age is 8.71. These two values remain relatively stable in the sample, alleviating the concern that our result would be driven by the time series variation of bond characteristics. Nevertheless, in our analysis we control for maturity of long and short maturity bonds. The average coupon is 2.97%, with a standard deviation of 1.82% and that maturity dates of bonds range between February 2005 and August 2046.

3.4.2 German sovereign bond market

The Federal Government of Germany is one of the largest issuer of government securities in the Eurozone. These securities are not only highly liquid; they also carry small issuer risk and managed to preserve their triple-A rating even during the Euro crisis. German government bonds have maturities ranging from 6 months to 30 years and they practically span the entire yield curve with 60-70 tradable securities at any point in time. Capital market securities consist of four types: Federal Treasury notes (Schtze), Federal notes (Bobls) and Federal bonds (Bunds) with the fore mentioned maturities ranging between 2 and 30 years. Since 2006, the Federal Government has also been issuing inflation-linked securities, but in this study we restrict our focus to nominal bonds.

German sovereign bonds are typically placed to primary dealers in the form of single issues via auctions, which can be followed by multiple re-openings. The average outstanding volume of a single issue is around 15-20 billion. Re-openings help to keep the market liquid and facilitate delivery of futures contracts on these bonds. The relative share of each security type is relatively stable over time: 2-year notes constitute 9%, 5-year notes 21%, 10 years Bunds 44,5% and 30 year Bunds 17% of the overall public debt issued by the Federal Government in 2015. This means that our sample of these bonds cover about 90% of all German tradable government debt. Moreover, these securities also account for 70% of the total issuance. All capital market securities have fixed maturities with redemption on maturity at the full nominal value, as well as annual fixed interest payments. These assets are also repo and ECB-eligible, moreover, Bunds are strippable.[8]

[8]We believe that this feature does not give rise to a large enough convenience yield in the respective bonds that this would distort our results.

3.4.3 Curve fitting

To fit the term structure of interest rates, the two most common approaches are the estimation of either spline-based or function-based models. The latter is a parsimonious parametric function that describes the entire yield curve. We choose to work with the model of Nelson and Siegel (1987), which provides a smooth, flexible parametric function. As De Pooter (2007) describes, this model is capable of capturing many of the typically observed shapes[9] that the yield curve can take over time. The Nelson-Siegel model assumes the following functional form for the instantaneous forward rate f:

$$f(m,b) = \beta_0 + \beta_1 \exp\left(-\frac{m}{\tau}\right) + \beta_2 \left(\frac{m}{\tau}\right) \exp\left(-\frac{m}{\tau}\right), \tag{3.1}$$

where m denotes time to maturity and $b = (\beta_0, \beta_1, \beta_2, \tau)$ are model parameters to be estimated. The three elements of the above sum have a clear interpretation: β_0 is the long term component, which is the same for every maturity; the component on β_1 is the short term component, which starts at 1 and decays to zero at an exponential rate; whereas the one on β_2 is the middle-term component, which can generate hump shapes of the yield curve. τ is a time invariant constant, which determines at which maturity the middle-term component reaches its maximum. Additionally, Diebold and Li (2006) show that the tree betas can be interpreted as three dynamic latent factors, governing the level, slope and curvature of the yield curve, respectively.

Following HPW, we use the parameterized forward curve of Equation 3.1 to derive the corresponding zero-coupon yield curve. This can be used to price any coupon-bearing bonds. Consequently, we can use market prices of bonds to back out the parameters in b. In particular, on each day of our sample we estimate the above model from observable market prices of German notes and bonds in our sample with remaining maturities between 6 months and 20 years. The result of this exercise on that day is a vector of parameters, based on which we can predict the theoretical values of the observed bonds. In order to fit the model, we choose parameters b_t by minimizing the weighted sum of squared deviations between the observable and the model implied prices:

$$b_t = \underset{b}{\text{argmin}} \sum_{i=1}^{N_t} \left[(P^i(b) - P_t^i) \times \frac{1}{D_i} \right]^2, \tag{3.2}$$

where N_t is the number of bonds for a given day, P_t are the prices observable on the market, whereas $P^i(b)$ is the model-implied price for bond i given the model parameters b, and D_i is the MaCaulay's duration for bond i. Following the yield curve fitting literature, we apply duration-weighting to bonds when estimating the parameters that minimize

[9] Which are monotonic, humped and S-shaped.

their pricing errors.

In fact, one deviation from the approach of HPW is that we do not estimate all parameters in Equation 3.1 freely, but keep τ fixed for the analysis. In preliminary unrestricted estimations τ turns out to be badly identified with very extreme values on days when the yield curve is flat. Due to suggested multicollinearity of the parameters in our sample, we decided to fix τ at its the median value (2.58) based on preliminary unrestricted estimation of the model. Then keeping τ fixed at its unconditional median, we re-estimated the rest of the parameters and obtained the fitted curves for each day in our sample.

Although we conduct most of our analysis keeping τ fixed at its unconditional median, we find that results are highly similar when we use a completely unrestricted estimation of tau allowing τ to be different day by day. Nevertheless, we added a robustness section, where we show how sensitive the resulting noise ad bias estimates are to different values of τ, e.g. half, twice or ten times the unconditional mean value.

Figure 3.1 depicts the fitted par-coupon curves and the observable yields for three days in our sample: January 3, 2005, September 15, 2008 and February 23, 2015. The left panels depict the short end of the yield curve, bonds with less than 20 years to maturity to which the curve is fitted; whereas the right panels show maturities longer than 20 years. For these bonds we extrapolate the fitted curve, thus the long end serves as an out-sample test of fit. The first day is a normal day', we can see that the short end on the curve fits rather well to the increasing term structure, while in the long end the deviations are larger, around 10-15 basis points. For the day of the Lehman bankruptcy, September 15, 2008, fitting errors on both the short and long ends are larger, so are they some days after the announcement of the quantitative easing of ECB on February 23, 2015.

3.4.4 Measures of noise

To construct the noise measure, we follow HPW. For each day in the sample, given a set of parameters b_t from Equation 3.2, they calculate the difference between the observable market yield and their model implied counterparts. They define noise as the dispersion of yields around the (entire) fitted yield curve, which is measured as the root mean squared pricing error in a given cross-section of bonds. Our approach differs from theirs in that we want to focus on the shorter and long ends of the curve separately. Consequently, we construct two sets of noise measures the following way.

For each date t, let b_t be a vector of parameters with $\tau = 2.58$. On the given date t, there are $N_{\text{short},t}$ notes and bonds in our sample with maturities between 6 month and 20 years and $N_{\text{long},t}$ bonds with remaining maturities exceeding 20 years. Then for each of these bonds part of $N_{\text{short},t}$ and $N_{\text{long},t}$, we calculate the difference between y_t^i and $y^i(b_t)$, the observed market and the model-implied yields, respectively. Then we define noise as dispersion around the given segment of the fitted curve, which we measure as root mean squared pricing error of the short and long ends of the yield curve separately:

$$Noise_{\text{short},t} = \sqrt{\frac{1}{N_{\text{short},t}} \sum_{i=1}^{N_{\text{short},t}} \left[y_t^{\text{short},i} - y^{\text{short},i}(b_t)\right]^2}; \tag{3.3}$$

$$Noise_{\text{long},t} = \sqrt{\frac{1}{N_{\text{long},t}} \sum_{i=1}^{N_{\text{long},t}} \left[y_t^{\text{long},i} - y^{\text{long},i}(b_t)\right]^2}. \tag{3.4}$$

To fit the yield curve, we restrict our sample to bonds with maturities between 6 months to 20 years, but we construct the noise measure for bonds with a wider range of maturities. Consequently, the noise measure of the longer end also serves as out-of-sample' test of the fitted curves. Our choice of the 20-year breakpoint is not arbitrary; we base it on the ultimate forward rate discussion of Solvency II. The ultimate forward rate is an extrapolation method that used to calculate long term discount rates used in valuation of liabilities for regulatory purposes for maturities beyond the last liquid point. This point for euro-denominated interest rates is set at 20 years to maturity. In addition, the yield curve beyond 20 years to maturity could be structurally distorted due to liquidity, market segmentation, demand pressure and differences in default premium across the curve. This is one of the questions we aim to answer in this paper.

Table 3.1 presents summary statistics for the noise measures, whereas Figures 3.2 and 3.3 depict the different noise measures. Panels A and B of Table 3.1 and Figure 3.2 focus on the noise measures of the short and long ends of the yield curve. The average short noise is 7.91 basis points between 2005 and 2015, whereas the long average is more than twice as much: with 19.39 basis points. In the long end, the standard deviation of noise is also larger. The range of noise is between a couple of basis points to 20.46 in the short and 49.58 in the long end. Figure 3.2 shows that the magnitude of the noise measures varies over time with large peaks around the financial and euro crises. Long yields have a poorer fit during this period, with their noise persistently staying high. However, this fit is still much better than what the regulatory discussion (on ultimate forward rates) considers as a starting point, especially given that we use such a simple curve fitting method.

Figure 3.3 compares our measures of noise to that of Hu et al. (2013). Panel C of Table 3.1 also gives descriptive statistics for the US noise measure. Its mean is 3.1 basis points with the range 0.72 to 20 basis points. The three measure exhibit a large spike at the financial crisis, but the US measure goes back to near zero afterwards, whereas the German measures, in reaction to the euro crisis, fluctuate at higher levels. However, we should point out that these three measures are not fully comparable due to their different composition and methodological choices. First, HPW fit their model to bonds of maturities shorter than 10 years, also including T-bills in their analysis. Noise is calculated for a somewhat smaller set of bonds and averaged over the whole curve. As opposed to this, we exclusively focus on notes and bonds. We fit the yield curve to bonds with maturities up to 20 years, but part of our analysis focuses on issues with even longer tenors. Noise is also separately calculated for short and longer maturity bonds. Second, due to our smaller cross-section of bonds we choose to fit a Nelson-Siegel curve, whereas HPW use the more elaborate Svensson method.

Despite the differences, we believe that our noise measures are proxies for market illiquidity, which show how prices deviate from fundamentals. To demonstrate this, the next section presents the link of noise measures to various asset level and market liquidity proxies. We analyze the direction and magnitude of the effects, especially focusing on the economic effect that is captured by one standard deviation change in explanatory variables or by 1% change in the asset growth of the ECB balance sheet.

3.4.5 Do German noise measures capture illiquidity?

Hu et al. (2013) show that the noise measure is a proxy for market-wide liquidity that is related to the arbitrage capital available in the market: the pricing error is larger and more persistent when funding liquidity conditions deteriorate and arbitrageurs are lacking resources to trade substantial price deviations away. To this extent, this measure captures funding liquidity, as in Brunnermeier and Pedersen (2009) or in Fontaine and Garcia (2012). We construct a pair of similar measures, but instead of aggregating all information along the yield curve, we focus on deviations of long and shorter maturities separately.

Table 3.2 reports the results of monthly regressions of changes (first differences) in short and long noise measures separately, regressed on a series of asset and market level liquidity proxies. Panel A shows that changes in the short end noise measure are linked to changing time to maturity or age of the bond. One standard deviation decrease in time to maturity, the bond gets 30 days older, increases the noise by 0.67 basis points. That is, that the older the bonds gets, the more likely it gets locked up in buy and hold investors' portfolios, which decreases the issue's liquidity, which in turn increases deviation from the theoretical yield. Similarly, short noise is linked to the changes in on-the-run spread of 10-year bonds and the Ted spread, with smaller but positive economic effects, which result in 0.29 and

0.38 basis points increases, respectively. This suggests that as the level of market liquidity or funding conditions worsen, noise tend to increase.

Panel B shows that the noise of long maturity bonds increases with the proportion of zero returns. If zero returns change by 0.63%, the noise increase 0.56 basis points. As for market liquidity proxies, long noise significantly links to US noise, which is a crisis-related aspect of funding liquidity. If the latter changes by one standard deviation, noise of German long maturity sovereign bonds increases by 0.91 basis points. Interestingly, the measures are not related to the implied bid-ask spread of Roll (1984) or to liquidity risk measured as the unexpected changes of the ILLIQ measure of German nominal bonds market. However, even the statistically significant connections are not very strong as the R^2 of these regressions are rather low, between 2 to 15%. HPW find similarly weak links to other measures of liquidity, which they claim to be an indication of noise capturing an aspect of liquidity beyond the typical measures used in the literature.

To summarize, we find that our measures of noise from the short and long ends of the yield curve capture somewhat different aspects of liquidity. Short noise is related to the age of the bond, as well as to market and funding liquidity; while long noise is linked to asset level illiquidity (zero returns) and the US noise measure. Nevertheless, similar to HPW, we find that even these statistically significant relationships have low explanatory power. This suggests that noise captures an aspect of illiquidity that correlates, but goes beyond conventional asset and market liquidity measures.

3.5 Segmentation effects: empirical evidence

3.5.1 The bias

To examine whether the long end of the German yield curve exhibits structural distortions that could result from market segmentation, we construct a measure, which is the average pricing error of long maturity bonds. This distortion we call the bias:

$$Bias_t = \frac{1}{N_{\text{long},t}} \sum_{i=1}^{N_{\text{long},t}} \left[y_t^{\text{long},i} - y^{\text{long},i}(b_t) \right]. \tag{3.5}$$

The bias is the average of differences between the observed, $y_t^{\text{long},i}$, and model implied yields, $y^{\text{long},i}(b_t)$, of bond issues with maturities exceeding 20 years.

Panel D of Table 3.1 reports the descriptive statistics of the bias. To complement this,

Figure 3.4 is composed of two panels. The upper panel depicts the overall bias, whereas the lower one differentiates between the maturity bucket of 20-25 years and maturities longer than 25 years. Table 3.1 suggests that the average bias throughout most the sample period is negative, -9.25 basis points, and ranges between -47 basis points to 25.74. Interestingly, it switches sign twice around the financial crisis, but stays persistently negative for the rest of the sample period. Figure 3.4 shows that the bias is not only negative after the financial crisis, but this effects is even more pronounced for the 25+ years to maturity bucket. Nevertheless, the figure does not indicate a clear trend which could result from, for instance regulation-induced demand pressure.

According to Equation 3.5, negative bias means that the observed yields are consistently lower than those implied by our pricing model. Could this be because those yields are distorted by demand pressure of clientele investors, or flight-to-safety, liquidity or credit risk premiums? To provide answer to this question, Table 3.3 presents evidence of the results of monthly regressions (of first differences) in which we link the bias to asset and market level liquidity measures, proxies for demand pressure, default risk, both in a uni-, and multivariate setting.

According to Panel A of Table 3.3, the link between the bias and bond level liquidity proxies, namely the Roll measure and time to maturity are both statistically and economically significant. If change in Roll's implied bid-ask spread increases by one standard deviation, thus bond liquidity decreases, the bias would get larger by 0.91 basis points. Similarly, if time to maturity decreases by 30 days, or alternatively if the bond gets a month older, the bias is also likely to grow by 0.75 basis points. The bias is also linked to changes in liquidity risk and flight-to-safety flows. Liquidity risk is measured as the surprise component of the monthly ILLIQ measure of the German nominal sovereign bond segment, while flight-to-safety flows are proxied by the PCA of CDS spreads of Eurozone periphery countries. We see that as credit or default risk of periphery countries increases (by one standard deviation), long maturity bond yields get 2.59 basis point further from the predicted smooth curve. Similarly, the bias seems to react to changes in German credit risk, as well as to changes in volatility or investor sentiment captured by the VIX index.

Panel B of Table 3.3 presents the results of multivariate analysis. Column 1 pools together all asset and market level liquidity proxies. The Roll measure is still significant with an even larger economic effects (1.33 basis points), while the effect of time to maturity and liquidity risks are washed away by the inclusion of the US noise measure. This latter implies that as US funding conditions worsen and pricing errors on Treasuries grow, the bias on German long maturity bonds also get wider. Column 2 looks at variables of demand pressure, default risk and volatility combined, out of which the German credit quality seems to matter the most. If the German CDS spread increases by 7.24 basis points, the bias also increases by 1.28 basis points. And finally, in Column 3 we pool all explanatory variables together and find that the Roll implied bid-ask spread stays significant with an effect size comparable to previous cases (1.31 bps), while the the Ted spread, our proxy for funding liquidity, is also significant, but with a sign opposing our expectations.

All in all, we find that the bias for most of the sample period is negative. This means that the observed yields are consistently below those implied by our model. In search for causes, we consider various proxies for demand pressure of clientele investors, flight-to-safety, liquidity and credit risk premiums. We find that the bias is linked to the Roll implied bid-ask spread, time to maturity of bonds, liquidity risk, flight-to-safety flows and the credit risk of Germany in a univariate setting. If considered in multivariate regressions, the Roll measure, German CDS and Ted spread are robust to the inclusion of other varables. In the next section we further analyze why long yields are too low' by looking at differences in drivers of yields of short and long maturity bonds.

3.5.2 Decomposing short and long maturity yields

If due to regulation, institutional features or given holding period, certain investors have limited access to the yield curve, this can create clientele demand, specific to certain bond maturities. Alongside with this, safe haven flows during the Euro crisis or large scale asset purchases of central banks and ECB can also put demand pressure on certain parts of the curve. Should these phenomena coexist with market frictions that limit arbitrage capital to flow into the sovereign bond market, the sum of these effects can persistently distort the part of the yield curve where it has the most prevalent effect.

Ideally one would like to test the influence of the above channels by formal cross-sectional pricing tests, but the data at hand, especially the small cross-section of long maturity bonds, does not grant us sufficient power for identification. Then the second best solution is to investigate the exposures of yields of short and long maturity bonds in the time series context. To do so, we decompose these yields in way similar to Krishnamurthy et al. (2015). According to their analysis, yields of Eurozone sovereigns contain the following distinguishable components: a risk free component, term premium, default risk premium, redenomination risk premium and a residual term that arises due to illiquidity frictions and segmentation. Taking their idea to our data, the following equation arises:

$$\begin{aligned} yield_{i,t} =& \frac{1}{T}\int_{s}^{s+T} \mathbb{E}[r_s]dt + Term\; pr._{\cdot i,t} + \\ &+ Liq.\; pr._{\cdot i,t} + Demand\; pr._{\cdot i,t} + Default\; pr._{\cdot i,t} + Controls_{i,t} \end{aligned} \tag{3.6}$$

where yields are explained by a risk free component based on the expectations hypothesis, a term premium and components that test segmentation of yields and liquidity with additional controls. The idea behind this decomposition is that it allows for separate examination of long and short yields, moreover, the resulting two sets of regressions coefficients are directly comparable.

To capture the first two components, we construct the risk free curve as the intersection of the EONIA overnight swap and euro swap curves. In our regressions we regress the average residual yields, the average of the difference between these two components and the observable yields from the market, on proxies of demand pressure, default premium, liquidity and other controls We proxy for demand pressure by including a time trend to capture regulatory changes; a variable capturing safe haven flows, derived from CDS of distressed Eurozone countries; and the percentage growth of the asset side of ECB's balance sheet, controlling for LSAP and the effects of QE. We also decompose the default premium component to the credit risk of the issuer, captured by the German CDS spread, and to breakup risk proxied by the KfW spread. This latter controls for selective default premium in which case the German government would default only on a set of its obligations, e.g. agency bonds or certain maturities. To show the robustness of our results, we also include additional controls, such as measures of market liquidity, we also control for the maturity structure by adding the time-series of the average time-to-maturity for daily cross-sections, as well as we add the US noise measure. We also examine whether there are liquidity spillovers between the short and long ends of the term structure by adding the respective noise measures to the regressions.

The resulting empirical relationships for the average short and long residual yields are the following:

$$\begin{aligned} \Delta[\overline{yield_{short,t}} - \overline{swap\; rate_t}] =& \beta_0 + \beta_1 \Delta Noise_t^{\text{short}} + \beta_2 Time\; trend + \beta_3 \Delta ECB\; Asset_t + \\ &+ \beta_4 \Delta FTS + \beta_5 \Delta Issuer\; credit\; risk + \beta_6 \Delta Breakup\; risk + \\ &+ \beta_7 \Delta Noise_t^{\text{long}} + B \cdot Controls_t + \varepsilon_t \end{aligned} \tag{3.7}$$

and

$$\begin{aligned}\Delta[\overline{yield_{long,t}} - \overline{swap\ rate_t}] =& \beta_0 + \beta_1 \Delta Noise_t^{\text{long}} + \beta_2 Time\ trend + \beta_3 \Delta ECB\ Asset_t + \\ &+ \beta_4 \Delta FTS + \beta_5 \Delta Issuer\ credit\ risk + \beta_6 \Delta Breakup\ risk + \\ &+ \beta_7 \Delta Noise_t^{\text{short}} + B \cdot Controls_t + \varepsilon_t \end{aligned} \tag{3.8}$$

Panels A and B of Table 3.4 report the results for Equations 3.7 and 3.8, respectively. Column 1 of both panels show the results without additional controls, while Column 2 includes all additional controls. At first glance, it seems that the noise measure is more important for long maturity bonds. Its effect not only survives the inclusion of other liquidity proxies and the US noise measure, but it is statistically and economically significant. One standard deviation change in the change of long noise, 4.23 basis points, leads to a shift of 1 yield basis point. The larger the pricing error, measured by the noise measure, the larger yields will be, which translate into lower prices due to illiquidity. Interestingly, the inclusion of the opposite noise measure does not affect yields on either ends of the curve: there is no evidence for liquidity spillover between the two segments.

If we look at the channels of demand pressure, we see that the time trend is insignificant with a negligible economic effect, whereas ECB asset purchase activity has an impact on shorter maturities. The effect of LSAPs is measured by the change of the natural logarithm of the asset side of ECB's balance sheet. It shows that 1% change in ECB asset growth decreases short yields by 0.28 basis points. This translates into an increase in bond prices, which is in line with decreased supply, providing suggestive evidence of ECB policies affecting sovereign yields. The third channel of demand pressure are flight-to-safety flows. We assume that the occurrence of these flows coincides with deteriorating credit quality of periphery and other large euro area countries. This has an effect on both short and long maturity bond yields to a similar extent. We see that one standard deviation decrease in periphery credit quality increases short and long yields by 3.02 and 2.94 basis points, respectively.

Next, we also look at whether default premium differs across short and long maturity bonds. We find that changes in the German CDS trigger a significant positive effect on yields: if the CDS changes by one standard deviation, short and long yields increase by 0.96 and 1.84 basis points, respectively. This means that as German credit quality deteriorates, yields will go up, while bond prices will decline. Nevertheless, this effect is rather small, taking into account that 1) Klinger and Lando (2015) show that safe haven CDS contracts are poor proxies for credit quality, as their demand is primarily driven by regulatory hedging requirements of investment banks 2) German CDS do increase over the euro crisis, but it stays low and stable relative to other Eurozone CDS contracts. We also examine how average yields correlate with breakup risk, as in De Santis (2015) or Simon (2015). We find that changes in the KfW spread, the yield spread on German agency and maturity-matched Treasury bonds, widens by 6.62 basis points, short yields decrease by 1.18, whereas their long counterparts by 1.42 basis points. The negative coefficient on the spread is inconsistent with the idea of breakup risk, it rather captures a flight-to-safety premium in sovereign bonds compared to the KfW agency bonds.

Interestingly, the constant terms of both specifications are insignificant, which suggests that the above factors not only capture the driving forces of yields, but also pick up their general downward trend in our sample. The R^2 of the decomposition is almost twice as high for short maturity bonds than for long ones. Moreover, all the fore mentioned effects are robust in sign and also in magnitude to the inclusion of additional controls, including the off-the-run spread, Ted spread and liquidity risk to control for market liquidity, the US noise measure and the average time to maturity of the given yield segment. Figure 3.5 depicts the rolling correlation of changes in short and long yields and provides further proof of negligible segmentation effects: we cannot observe any major trends in the comovement of these yields. Their correlation has a noticeable drop around the financial crisis, but even at its bottom it is 65%, while it peaks in 2012 during the euro crisis at 93.5

Overall, the presented yield decomposition reveals little evidence of substantial segmentation in German sovereign bond markets. Short yields are mostly driven by short noise and large scale asset purchases, while both long and short yields are strongly linked to flight-to-safety flows and to German credit quality. Our results show that although there are statistically significant differences in the drivers of short and long maturity bonds, long yields are not substantially distorted by demand pressure, default or liquidity premiums.

3.5.3 Yields and liquidity

This section studies the link between yields and different market wide and maturity-specific liquidity measures. Our aim is to discover whether liquidity has a differential effect on short and long maturity bond yields. Although our bond sample does not allow for formal cross-sectional pricing tests of liquidity, to answer this question we investigate the exposures of yields to the different liquidity measures in a time series context. Moreover, we also allow for liquidity spillovers between the two ends of the yield curve and scrutinize how the financial crisis, a major liquidity dry up, affected liquidity characteristics of German sovereign bonds.

Table 3.5 reports results of yield differences regressed on a set of segment specific and market wide liquidity measures. Panel A shows the results of bonds with less than 20 years to maturity. The average short yield is significantly related to changes in the corresponding noise measure: if short noise increases by 1.75 basis points (one std.), the average short yield decreases by 0.92 basis points. This effect is robust to the inclusion of a wide range of market liquidity measures, especially that of the Ted spread, although previous analysis suggested that the noise measure is also related to funding and market liquidity. Additionally, the Ted spread is also highly significant, with an economic effect of 1.71 bps yield change if the spread widens by 25 basis points; alongside with a large increase in the R^2 from 5% to 26% if market liquidity measures are included. Another feature of the short noise measure is that its effect on yield varies over time as can be seen if it is interacted with an indicator variable of the financial crisis. The sign and magnitude of the coefficient of noise changes between the pre and post crisis and no crisis regimes: after the financial crisis changes in the short noise measure decrease average short yields (by 1.17 bps), while before it has a less significant, but positive coefficient, which doubles in size compared to the crisis period. Usually one would expect that yields go up when liquidity decreases, thus finding contradictory evidence over the crisis is puzzling.

Panel B focuses on liquidity of long maturity bonds. Similarly to the average short yield, the average long yield has a tie to the fitting error of long maturity bonds. This relationship is definitely weaker than that of short yields and short noise, and is not robust to the inclusion of market liquidity measures. However, the economic effect is comparable: 0.89 basis points. Besides, unlike for the average short yield, there is evidence for a liquidity spillover from the short end to the long one: if short noise increases by one standard deviation, the average long yield would increase by 0.75 basis points. Interestingly the difference between the pre and post financial crisis periods is smaller that of short bonds: after the crisis the long noise measure does not have a statistically significant relation with the average long yield, whereas the interaction of long noise and the pre-crisis period is borderline significant with a positive coefficient. R^2 of all five specifications are low, in the range of 3 to 7

In conclusion, we find that yields of the short and long ends of the yield curve are linked to a wide range of liquidity measures. The effect of short noise on the average short yield is statistically and economically significant and robust to the inclusion of other measures. Moreover, the effect of short noise varies over time in relation to the financial crisis. As opposed to this, the average long yield has a weaker connection to long noise measure, while we find evidence of liquidity spillover from the short to the long end of the yield curve. The next section further discovers variables that could be driving the two noise measures.

3.5.4 Drivers of liquidity segmentation

We have not found evidence for substantial segmentation in yields, but the differential effect of long and short noise measures suggests that liquidity of the corresponding bonds is less correlated than their yields are. This low correlation could be indicative evidence of liquidity segmentation. In section we provide further evidence, by exploring whether these differences stem from the underlying forces that drive liquidity of each segment.

Table 3.6 compares how short and long noise measures relate to demand pressure, default risk and volatility measures in Panels A and B, respectively. In previous analyses we linked the noise measures to various asset and market liquidity proxies. We concluded that the short noise measure captures market liquidity, which is beyond the off-the-run and Ted spreads; while the long noise seems to behave independently from other market variables. As expected, there are also some differences in their exposures to the fore mentioned measures: short noise significantly correlates with flight-to-safety flows and the asset growth of the ECB's balance sheet. As the credit quality of periphery countries worsen, the change in short noise increases by 0.33 basis points. This is due to the increased demand in safe haven bonds, which pushes prices away from the theoretical value suggested by the fitted yield curve. Similarly, as ECB purchases shorter maturity German bonds, 1% growth of the ECB assets increases noise by 0.57 basis points. Panel A also suggests that noise is sensitive to volatility captured by the VIX index or the German stock market volatility index, DAX.

Panel B shows that long noise has a similar relationship to safe haven flows as short noise. This suggests that these flows are not exclusive to certain maturities, hence the overall German yield curve becomes attractive if periphery credit quality declines. Interestingly, long noise is more exposed to changes in German credit quality or breakup risk, captured by changes in the German CDS and the KfW spread, respectively. Their economic effects are also sizable: if the CDS spread increases by 7.28 basis points, the noise on the long end of the curve increases by 1.04 basis points; while 6.62 basis point change in the KfW spread triggers a similarly sized: 1.19 basis point increase. Long noise is also correlated with volatility, with economic effects that are twice as large as those of the short noise.

Figure 3.6 depicts the 180-day rolling correlation of changes in the short and noise measure. Noise measures exhibit high correlation before the financial crisis and during the post-crisis recovery period. There is a significant drop at the crisis, although the correlation stays positive, unlike after 2011, where the correlation becomes negative and persistently stays there for months. Similarly to the correlation between short and long maturity bond yields, we cannot see a clear pattern that would suggest a high level of liquidity segmentation, however this correlation has a considerably larger range: -45 to

90

In summary, we find that short and long noise measures are related to somewhat different market forces: short noise is sensitive to large scale asset purchases of the ECB, as well as to changes in stock market volatility and flight-to-safety flows. The long noise measure has a similar relationship to safe haven flows, which suggests that these flows are not exclusive to certain maturities, hence the overall German yield curve becomes attractive if periphery credit quality declines. The long noise is also sensitive to declining credit quality of the issuer, as well as to breakup risk, however all the above effects are economically rather small.

3.5.5 Robustness tests

In this section we aim to test how sensitive the long maturity noise and bias measures are to changing the shape parameter, τ, of the Nelson-Siegel curve. In preliminary unrestricted estimations, τ turns out to be poorly identified with very extreme values on days when the yield curve is flat. Consequently, we fixed τ at its the median value (2.58) and conducted most of our analysis keeping it fixed. In unreported regressions we find that results are highly similar when we use a completely unrestricted estimation of tau allowing tau to be different day by day.

Alternatively, we could also impose structure on the shape parameter, as Quaedvlieg and Schotman (2016) propose in their paper. Any moving average or ARMA process could be applied, as long as there are not prolonged periods of days with very flat term structure.However, imposing such a structure on the shape parameter is non-trivial. On the one hand, in our sample, especially around the onset of the financial and the peak of the euro crises, we have a multitude of days with very flat term structure. On the other hand, another obstacle for implementing their procedure is that it is numerically challenging when using coupon bonds, since in this case one cannot filter out the other parameters on each day from the zero yields similarly to Quaedvlieg and Schotman (2016).

Given our data limitations, we choose a more feasible alternative, which is to show that the size and time series characteristics of the long noise measure and the bias do not depend on the choice of the shape parameter, τ.[10] The four panels of Figure3.7 aim to prove this point: they depict the noise and bias measures for bonds with maturities longer than 20 years, from unrestricted estimation or τ fixed at different values.

[10] We focus on measures of the long end as those are 'out-of-sample' tests of the curve, where deviations and thus the differences arising due to alternative values of τ are expect to be larger.

First, we compare the long noise and bias measures for τs from the unrestricted estimation and the one, where τ is fixed at its unconditional median. The upper panels of Figure3.7 show that in both cases, maybe even more so for the bias, the time-series patterns are rather similar. The most significant difference is that the series based on the unrestricted estimation exhibit more spikes, those are typically days, where the numerical optimization did not converge to global maximum. To circumvent this problem, we decided to restrict the shape parameter at its median value based on the unrestricted results.

How much our results are driven by the choice of τ? In answering this question, we estimate the long noise and bias measures given different fixed levels of τ. These are half of the median, median, twice and ten times the unconditional median. Since τ is the time invariant constant, which determines at which maturity the middle-term component reaches its maximum, we believe that our choice or at most twice the median could be reasonable values. The lower panels of Figure3.7 show that even an unrealistically large τ, 25.8, would not generate noise and bias measures that are inexplicably large. While the median value produces an average long noise and bias of 18.82 and -9.25 basis points, taking this extreme value leads to 50.26 and -14.77 basis points, respectively. These values still fall in the reasonable range, despite that such a τ economically would not make sense.

3.6 Discussion and policy implication

This section aims to address three issues: first, we would like to compare the extrapolated Nelson-Siegel curve to the one used by the regulator for liability valuation of pension and insurance funds. Second, we perform a thought experiment, in which we quantify economic effects from the asset and liability management point of view of two hypothetical pension funds with different characteristics. And lastly, we explain the European policy discussion and specify our contribution thereof, also pointing towards directions for further research.

3.6.1 Regulatory vs. extrapolated Nelson-Siegel curves

How different is the extrapolated Nelson-Siegel curve from the one used by regulators? In answering this question, we compare the ultimate forward rate (UFR) curve provided to us by the Dutch Central Bank (DNB) to the one we propose in this paper. The UFR method is an extrapolation technique to calculate long term discount rates for valuation of liabilities for regulatory purposes, for maturities beyond the last liquid point. This point for euro-denominated interest rates is set at 20 years and the curve reaches its maximum value of 4.2% well beyond 60 years. The UFR curve provided by DNB is constructed followings the guidelines provided by EIOPA and it is based on German interest swap data from Bloomberg.

Figure3.8 presents the comparison between the UFR and the extrapolated curve we fitted on German bond data. The upper panel depicts the UFR curve as it has been provided by the DNB. It is immediately apparent that fitting the curve on swap rates gives rise to an approximate 35 basis point differential between the two term structures. This is most likely the sum of compensation for counterparty risk in bilateral swap transactions and illiquidity of different contracts maturities. Nevertheless, this differential seems to widen especially beyond the last liquid point, where our market-based estimates for long term yields are much lower than their swap-based UFR counterparts. To confirm this increase in the difference, the lower panel shows the same Nelson Siegel curve together with the UFR curve net of swap premium. The assumption underlying the swap premium correction is that the counterparty and liquidity risks are independent of the contracts maturity. Then we define the swap premium as the average difference between the NS curve fitted on bond data and the UFR curve based on swap data up to 20 years to maturity. The corrected curve is parallel to the original but then shifted downwards so that short yields are fairly similar and deviations for longer maturities reflect differences in the extrapolation.

This difference in extrapolation is likely to have a sizeable effect on liability valuation, which the next section will examine in more detail.

3.6.2 Liability valuation: a thought experiment

In previous sections we claim that the economic effects of segmentation are rather small: they are in the ballpark of 1-3 basis points. Eventually, by the use of leverage arbitrageurs could profit from any mispricing or from small risk premium differences between different maturities. However, a more interesting aspect of economic significance is how liability valuation of pension funds would be affected by switching from the UFR curve to our proposed term structure based on bond market data.

Our aim is to quantify the effect of different discount curves on liability valuation. To perform this thought experiment, we use the two curves on the lower panel of Figure 3.8. We assume there are two hypothetical pension funds. Both funds pay out €100 in total over a payment schedule of 60 years. Participants join a fund at the age of 25 and pay a steady stream of equal contributions over the years, until they reach retirement age at 65.

At 65 they get a steady stream of defined cash flows until they die at the age of 85. The first fund is a large fund with a constant number of participants uniformly distributed between the age of 25 and 85 years. This implies that in each period, the fund has to pay 1/60 fraction of its total liabilities, thus €1.67 out. The other fund is a young fund, in which participant are between the age of 25 and 50 years old and are in a constant supply. This means that in the first 15 years the fund does not have any payouts, but as participants gradually reach retirement age, payouts increase. They do so until the 35$^{\text{th}}$ year of the schedule, when the younger fund also reaches a steady state: equal number of people entering the fund at 25 and dying at the age of 85.

The upper panel of Figure 3.9 depicts the payout schedule of the two funds, while its lower counterpart shows the comparison between the difference in liability values discounted by either the UFR or the extrapolated NS curve.

We find that the choice of the discount curve matters: the present value of liabilities is smaller when the UFR approach is applied. Despite that the value difference is more sizeable for the younger fund, where the duration of the liabilities is longer, its magnitude is around 1% As opposed to this, the discount rates predicted by the extrapolated Nelson-Siegel curve result in higher liability values. Although, this analysis is incomplete without considering the asset side of a pension portfolio and the funding ratio, we refrain from further investigating the asset side given the large number of assumptions required to quantify these effects. Even by means of such simple analysis we can conclude that the choice of the discount rate matters: we find that by switching from one curve to another, the present values of liabilities are different.

Looking at the actual size of the difference and the volatility of the present value, we would have expected switching from the one curve to another to have a larger effect. This implies that the impact on the funding ratio of a typical pension fund is rather small, especially compared to the annual fluctuations of funding ratios due to market returns and interest rate changes. Note, however, that the presented difference is likely conservative as it relies on the UFR curve net of the swap premium, therefore the actual difference is likely to be more significant than presented here. By taking the swap premium out, we eliminated a potential distortion that might affect the overall UFR curve, without respect of maturity segments. Nevertheless, this thought experiment already points to a direct implication on asset and liability management: all else equal, larger liability value decreases the funding ratio. This smaller value on one side of the balance sheet would have to be compensated with larger asset value and/or potentially stricter risk management.

3.6.3 Policy discussion and the scope of our contribution

The primary aim of this paper is to explore yield and liquidity segmentation of (a representative) European nominal term structure. In particular, we analyze whether the behavior and drivers of short and long maturity bonds yields differ and how their liquidity characteristics are driven by investor groups with different holding periods. However, in doing so, this paper also contributes to the European policy discussion on the asset and liability management of pension and insurance funds. Namely, we suggest a viable alternative to the UFR method proposed by Solvency II and EIOPA. The UFR method aims to specify the construction of long term discount rates, by means of extrapolating swap rates beyond the last liquid point of the term structure.

There are three distinct issues regarding the current industry practice: 1) how to fit the curve and how to extrapolate beyond the last liquid point; 2) where to set the last liquid point; and 3) based on which information should long term discount rates be determined. This paper does not tackle all three of these issues. We propose a simple method to fit a smooth forward curve and to extrapolate that beyond the last liquid point. We only partially assess which extrapolation method, our simple extrapolation based on bond data or the UFR approach, performs better in the thought experiment of the previous section. Second, we take the last liquid point suggested by EIOPA and set it to 20 years to maturity. There is an open debate on where this point should be set, while EIOPA and the Netherlands uses 20 years for the division point, in Sweden the last liquid point is set at 15 years to maturity. Then the question arises, where is the actual segmentation point along the yield curve? And finally, our results suggest that long maturity bond yields might be appropriate for the valuation of long-term liabilities - especially if the distortion in long yields due to segmentation is smaller than risk premiums in observable swap quotes.

Nevertheless, finding that long maturity bond yields are not substantially distorted has important policy implications. Part of the policy discussion on valuation of pension and insurance liabilities for regulatory purposes is how to model long term discount rates. The current approach is based on the ultimate forward rate method, which in light of our results might seem unnecessary if yields beyond the last liquid are not distorted. This case one could extrapolate long term discount rates from yields observed in bonds markets. Then the question arises: is the ultimate forward rate discussion much ado about nothing? One the one hand, the answer is no, as there is a lot of money at stake: pension funds and insurers together managed €3.5 trillion worth of assets in 2015, thus making sure that this wealth is properly valuated is crucial. On the other hand, the answer could be yes: it seems that we can trust yields of long maturity bonds as any existing effects of yield or liquidity segmentation are quite small.

3.7 Conclusion

This paper examines the differential pricing of short and long maturity bonds. Our aim is to study whether there is segmentation along the German yield curve and discover the channels through which this would affect the pricing of short and long maturity bonds. In summary, we find low levels of yield segmentation: our results suggest that although there are statistically significant differences in the pricing and drivers of short and long maturity bonds, the corresponding economic effects are rather small. This means that long yields are not substantially distorted by demand pressure, default or liquidity premiums.

Our empirical strategy is as follows. First, to study segmentation of long and shorter maturity bonds, we construct a pair of noise measures similar to that of Hu et al. (2013). The noise measures are liquidity proxies that are based on the pricing errors of observable yields relative to a smooth curve. We find that these measures capture different aspects of liquidity: short noise is related to the age of the bond and to market and funding liquidity; while long noise is linked to bond issue level illiquidity and the US noise measure. Nevertheless, conventional liquidity proxies do not fully explain the two noise measures suggesting that they capture an aspect of illiquidity that correlates but goes beyond those measures.

Next, to further examine the pricing of bonds, we also look at the average fitting error, called the bias. We find that the bias is persistently negative, meaning that the observed yields are consistently below those implied by our pricing model. To explain why long yields are too low', we link the bias to the Roll implied bid-ask spread, time to maturity of bonds, liquidity risk and the Ted spread, flight-to-safety flows and to the credit risk of Germany. To deepen our understanding of time-series drivers of short and long maturity bond yields, we study their exposures to proxies of demand pressure, default premiums and liquidity. The yield decomposition reveals little evidence of significant segmentation in German sovereign bond markets. Short yields are mostly driven by short noise and large scale asset purchases, while both long and short yields are strongly linked flight-to-safety flows and to German credit quality.

Looking at liquidity segmentation, we study how yields of the short and long ends of the yield curve are linked to a wide range of liquidity measures. We find that short noise has a significant and robust effect on average short yield, which varies over time. As opposed to this, we find that no such effect in the average long yield, while there is a liquidity spillover from the short to the long end of the yield curve. We also consider if short and long noise are related to different market forces and find that short noise is sensitive to large scale asset purchases of the ECB, and to changes in stock market volatility and flight-to-safety flows. We conclude that these findings are in line with the theory of liquidity segmentation (Amihud and Mendelson, 1986; Beber et al., 2012), the economic effects of liquidity segmentation are rather small. Nevertheless, we show that the nature of liquidity varies along the curve: liquidity of short maturity bonds seems more systematic in nature, whilst liquidity of long maturity bonds behaves independently from other market variables.

Finding that long maturity bond yields are not substantially distorted has important policy implications. Part of the policy discussion on valuation of pension and insurance liabilities for regulatory purposes is how to model long term discount rates. In light of our results the current regulatory practice and the use of the UFR method seems unnecessary if yields beyond the last liquid are not distorted. This case one could extrapolate long term discount rates from yields observed in bonds markets. Then the question arises: is the ultimate forward rate discussion much ado about nothing? One the one hand, the answer is no, as there is a lot of money at stake: pension funds and insurers together managed €3.5 trillion worth of assets in 2015, thus making sure that this wealth is properly valuated is crucial. On the other hand, the answer is yes: it seems that we can trust yields of long maturity bonds as any existing effects of yield or liquidity segmentation are quite small.

Table 3.1
Summary statistics of noise measures and the bias

The table presents descriptive statistics of different noise measures and the bias. Noise measures and the bias are derived from the difference between the observable market yields and their theoretical counterparts, based on Nelson-Siegel curves fitted on German bonds. Panels A and B contain information on short and long noise measures, respectively. Noise is the root mean squared pricing error of yields. Panel C present the noise measure of Hu et al. (2013) (HPW), which is based on US Treasuries, obtained from the website of Jun Pan. Panel D presents the bias, the mean pricing error of long maturity bond yields. All measures are in basis points.

Panel A. Short noise

	# of bonds	Mean	St.dev	Min	Max
2000-2015	38-56	6.4168	2.7791	2.0036	23.8380
2005-2015	38-55	7.9122	7.8396	2.3838	20.4675

Panel B. Long noise

	# of bonds	Mean	St.dev	Min	Max
2000-2015	1-8	18.8233	10.7720	1.7053	50.3190
2005-2015	1-8	19.3881	10.4078	1.7053	49.5880

Panel C. US noise measure from HPW (2013)

	# of bonds	Mean	St.dev	Min	Max
2000-Dec 2014	appr. 165	3.1009	2.6507	0.7232	20.4675
2005-Dec 2014	appr. 165	3.0891	3.1842	0.7232	20.4675

Panel D. Bias

	# of bonds	Mean	St.dev	Min	Max
2000-2015	1-8	-1.5346	19.9604	-47.9043	50.2931
2005-2015	1-8	-9.2537	17.4175	-47.9043	25.7488

Table 3.2
Monthly changes in noise measures regressed on other variables

This table reports OLS coefficients of monthly regressions in differences, where stars denote conventional significance levels. T-statistics are based on Newey-West errors with 3 month-lag. Panel A shows results for changes in short noise regressed on asset and market level liquidity proxies, whereas Panel B repeats the analysis for bonds with maturities longer than 20 years. Zero returns show the % of days with zero return, Roll is the return autocorrelation implied bid-ask spread, and time to maturity is measured in years. The US noise measure is from Hu et al. (2013). The on-the-run spread is the yield difference of newly issued and seasoned 10-year Bunds, liquidity risk is measured as the AR(1) residual of the aggregate ILLIQ measure of the German nominal bond market. Ted spread is the difference between the 3-month LIBOR and the US Treasure bill. The sample period is January 2005 - June 2015. All spreads and noise measures are in basis points. *, **, and *** denote statistical significance of the reported coefficients at the 10%, 5%, and 1% level, respectively.

Panel A. ΔShort noise

Asset liquidity

	(1)	(2)	(3)	(4)
ΔZeroReturn	-0.005			-0.078
ΔRoll		-0.169		-0.378
ΔTtm			-8.000***	-8.016**
R^2	0.00	0.00	0.15	0.15
# of months	124	124	124	124

Market liquidity

	(1)	(2)	(3)	(4)	(5)
ΔUS noise	0.253				0.186
ΔOntherun		0.193**			0.250***
ΔLiq. Risk			-2.710		-4.481
ΔTed				0.015**	0.011*
R^2		0.03	0.00	0.05	0.15
# of months		120	93	124	92

Panel B. ΔLong noise

Asset liquidity

	(1)	(2)	(3)	(4)
ΔZeroReturn	0.893**			0.922**
ΔRoll		-0.334		-0.584
ΔTtm			-1.876	-1.742
R^2	0.02	0.00	0.01	0.02
# of months	124	124	124	124

Market liquidity

	(1)	(2)	(3)	(4)	(5)
ΔUS noise	0.980**				1.151***
ΔOntherun		-0.178			-0.016
ΔLiq. Risk			6.644		2.928
ΔTed				0.000	-0.015
R^2	0.05	0	0	0	0.03
# of months	119	120	93	124	92

Table 3.3
Drivers of the bias – structural distortions in the curve

This table presents OLS coefficients of monthly regressions in differences, where stars denote conventional significance levels. T-statistics are based on Newey-West errors of 3 month lags. Panel A shows results of univariate regressions of the bias on liquidity, demand pressure, default risk and volatility proxies, whereas Panel B repeats the analysis in a multivariate setting. Variable description of liquidity proxies can be found of Table 3.2. Flight-to-safety is the first PCA of CDS spreads of distressed Eurozone countries, and ECB asset growth is the change in the logarithm of the asset side of the ECB balance sheet. KfW spread is the yield difference between 10-year KFW agency bond and a maturity-matched German Bund, and VIX is the volatility index issued by CBOE. The sample period is January 2005 - June 2015. All spreads and noise measures are in basis points. *, **, and *** denote statistical significance of the reported coefficients at the 10%, 5%, and 1% level, respectively.

Panel A: Univariate regressions

Asset liquidity

	(1)	(2)	(3)
ΔZeroReturn	-0.278		
ΔRoll		-2.560*	
ΔTtm			-3.131*
R^2	0	0.04	0.03
# of months	124	124	124

Market liquidity

	(1)	(2)	(3)
ΔOntherun	0.333		
ΔLiq. risk		-26.516**	
ΔTed			0.016
R^2	0.01	0.05	0.01
# of months	120	93	124

Demand pressure

	(1)	(2)	(3)
Time trend	0.004		
ΔFTS		-7.787***	
ΔECB asset			-0.109
R^2	0.00	0.08	0.01
# of months	124	124	124

Default and volatility

	(1)	(2)	(3)
ΔCDS	-0.224***		
ΔKfW		-0.056	
ΔVIX			-0.163*
R^2	0.12	0.01	0.02
# of months	124	101	124

Panel B: Multivariate regressions

	(1)	(2)	(3)
ΔZeroReturn	0.409		0.090
ΔRoll	-3.749**		-3.686**
ΔTtm	-2.374		-1.875
ΔUS noise	-1.283***		-0.498
ΔOntherun	0.291		0.401
ΔLiq. risk	-17.381		-13.117
ΔTed	0.022		0.035*
Time trend		-0.007	-0.006
ΔFTS		-2.602	-2.352
ΔECB asset		0.017	-0.117
ΔCDS		-0.176**	-0.142
ΔKfW		0.050	0.029
ΔVIX		-0.082	-0.091
R^2	0.20	0.15	0.31
# of months	92	101	92

Table 3.4
Yield decomposition - the effect of segmentation

This table reports results of monthly OLS regressions of yield differences. Variables with delta are differences, whereas time trend and the ECB asset growth are in levels. Short noise is the squared mean pricing error of these bonds relative to the smooth Nelson-Siegel curve. Description of variables can be found in Tables 3.2 and 3.3. Crisis dummy is defined to be 1 for the period after September 2008, whereas No crisis dummy equals 1 before that date. The sample period is January 2005 - June 2015. All spreads and noise measures are in basis points. The absolute value of Newey-West 3 month-lagged t-statistics are in parentheses, while *, **, and *** denote significance at the 10%, 5%, and 1% level.

Panel A. Decomposition of short yields

	(1)	(2)
ΔShort noise	-0.3214	-0.2692
	(1.59)	(1.29)
Time trend	-0.0111	-0.005
	(0.89)	(0.21)
ECB asset growth	-0.2807	-0.1472
	(2.84)***	(1.23)
ΔFlight-to-safety	-9.0968	-11.2594
	(2.88)***	(3.41)***
ΔCDS	0.1325	0.1779
	(1.91)*	(2.40)**
ΔKfW	-0.1787	-0.1583
	(2.89)***	(2.40)**
ΔLong noise	0.0713	0.0522
	(0.79)	(0.58)
Constant	7.0829	5.2419
	(0.92)	(0.55)
Time to maturity	No	Yes
ΔUS noise	No	Yes
Market liquidity pr.	No	Yes
R^2	0.31	0.46
N	101	92

Panel B. Decomposition of long yields

	(1)	(2)
ΔLong noise	0.2211	0.2333
	(1.73)*	(1.72)*
Time trend	-0.0162	-0.0255
	(0.90)	(0.72)
ECB asset growth	0.0752	-0.0755
	(0.54)	(0.42)
ΔFlight-to-safety	-8.8555	-8.4888
	(1.99)**	(1.71)*
ΔCDS	0.2538	0.2412
	(2.60)**	(2.14)**
ΔKfW	-0.2181	-0.2752
	(2.50)**	(2.73)***
ΔShort noise	0.3815	0.4264
	(1.34)	(1.36)
Constant	9.8386	21.3583
	(0.91)	(1.42)
Time to maturity	No	Yes
ΔUS noise	No	Yes
Market liquidity pr.	No	Yes
R^2	0.17	0.24
N	101	92

Table 3.5
Yields and liquidity

This table reports results of monthly OLS regressions of yield differences on different measures of liquidity. Panel A reports the results of yield decomposition for short maturity bond yields, whereas Panel B repeats the analysis for yields of bonds with more than 20 years to maturity. Variables with delta are differences, and all spreads and noise measures are in basis points. Noise measures are the squared mean pricing error of these bonds relative to the fitted Nelson-Siegel curve. Crisis dummy is defined to be 1 for the period after September 2008, whereas No crisis dummy equals 1 before that date. Description of liquidity measures can be found in Table 3.2. The sample period is January 2005 - June 2015. Absolute values of 3-month lagged Newey-West t-statistics are in parentheses, while *, **, and *** denote significance at the 10%, 5%, and 1% level.

Panel A. Liquidity of short maturity bonds

	(1)	(2)	(3)	(4)	(5)
ΔShort noise	-0.5216	-0.5236			-0.3834
	(2.65)***	(2.65)***			(1.74)*
ΔLong noise		-0.0275			-0.0611
		(0.34)			(0.63)
Crisis*ΔS.noise			-0.665		
			(3.30)***		
NoCrisis*ΔS.noise				1.2641	
				(1.72)*	
ΔTed spread					-0.0668
					(4.31)***
ΔLiq. Risk					-4.1554
					(0.39)
ΔOff-the-run spr.					0.2180
					(0.73)
ΔUS noise					-0.1218
					(0.28)
Constant	-0.0371	-0.0374	-0.0652	-0.1037	0.1011
	(0.11)	(0.11)	(0.19)	(0.3)	(0.24)
R^2	0.05	0.06	0.08	0.02	0.26
N	124	124	124	124	92

Panel B. Liquidity of long maturity bonds

	(1)	(2)	(3)	(4)	(5)
ΔLong noise	0.2124	0.2178			0.1431
	(1.98)*	(2.04)**			(1.04)
ΔShort noise		0.4319			0.4313
		(1.68)*			(1.37)
Crisis*ΔL.noise			0.1526		
			(1.20)		
NoCrisis*ΔL.noise				0.3726	
				(1.81)*	
ΔTed spread					0.0084
					(0.38)
ΔLiq. Risk					-1.1098
					(0.07)
ΔOff-the-run spr.					-0.4214
					(0.98)
ΔUS noise					0.6230
					(0.99)
Constant	-0.0038	-0.0109	-0.0004	-0.0165	0.3545
	(0.01)	(0.02)	(0.00)	(0.04)	(0.60)
R^2	0.03	0.05	0.01	0.03	0.07
N	124	124	124	124	92

Table 3.6
Drivers of bond liquidity

This table reports OLS coefficients of monthly regressions in differences, where stars denote conventional significance levels. T-statistics are based on Newey-West errors of 3-month lags. In Panel A short noise is regressed on demand pressure, volatility and default risk proxies, whereas Panel B repeats the analysis for bonds with maturities longer than 20 years. Description of variables can be found in Table 3.3. The sample period is January 2005 - June 2015. All spreads and noise measures are in basis points.*, **, and *** denote statistical significance of the reported coefficients at the 10%, 5%, and 1% level, respectively.

Panel A. ΔShort noise

Asset liquidity

	(1)	(2)	(3)	(4)
Time trend	-0.001			0.002
ΔFTS		0.993**		0.706
ECB asset gr.			0.120***	0.106**
R^2	0.00	0.04	0.07	0.08
# of months	124	124	124	124

Market liquidity

	(1)	(2)	(3)	(4)	(5)
ΔCDS	0.020				0.022
ΔKfW		-0.026			-0.051*
ΔVIX			0.071**		0.278**
ΔDAX				0.005	-0.239*
R^2	0.01	0.01	0.03	0.01	0.09
# of months	124	101	124	114	99

Panel B. ΔLong noise

Asset liquidity

	(1)	(2)	(3)	(4)
Time trend	-0.011			-0.001
ΔFTS		2.979***		0.282**
ECB asset gr.			0.098	0.002
R^2	0.01	0.05	0.01	0.06
# of months	124	124	124	124

Market liquidity

	(1)	(2)	(3)	(4)	(5)
ΔCDS	0.143***				0.088
ΔKfW		0.181***			0.131*
ΔVIX			0.151*		-0.064
ΔDAX				0.019*	0.178
R^2	0.06	0.08	0.02	0.03	0.12
# of months	124	101	124	114	99

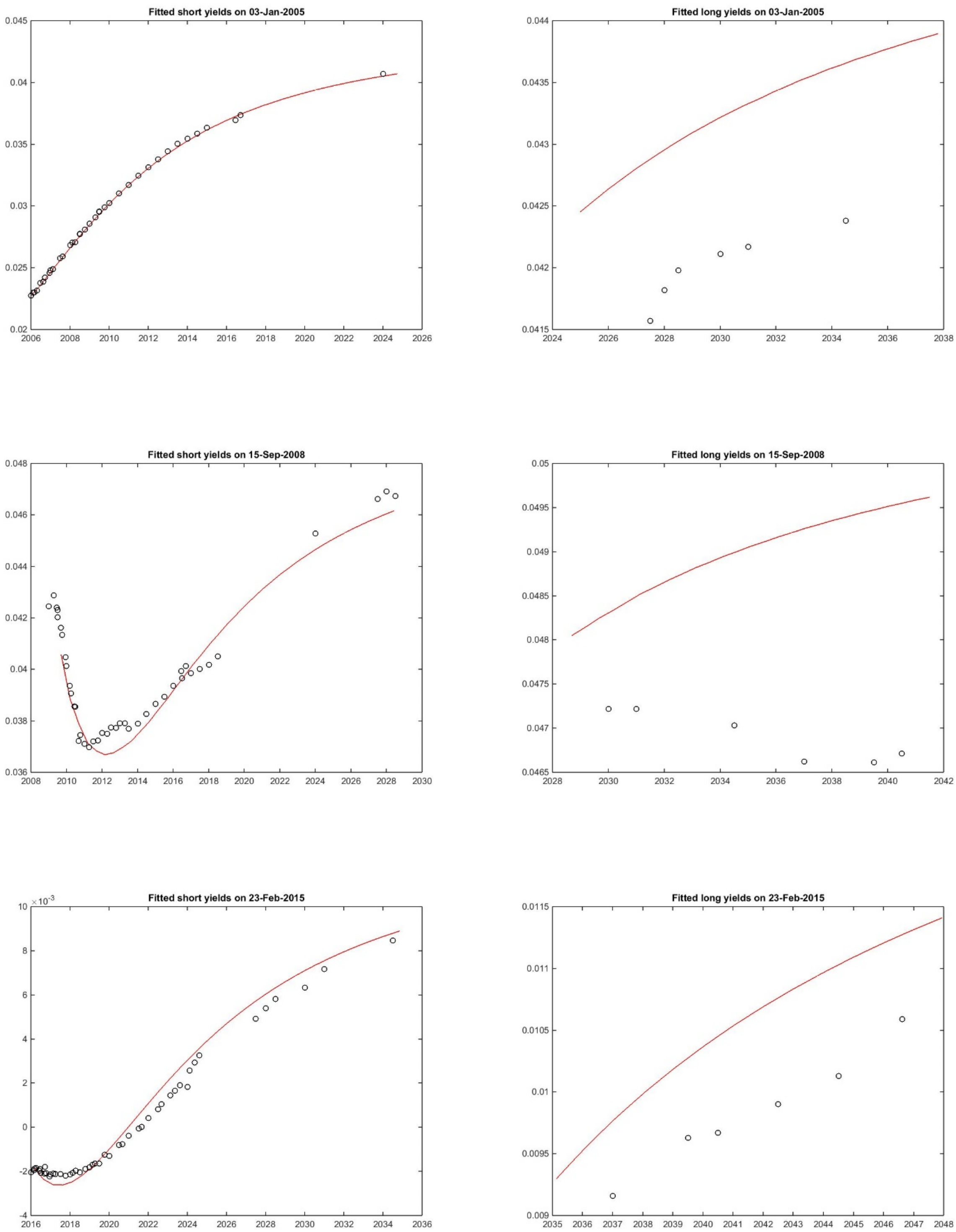

Figure 3.1 Fitted par-coupon yield curves and observable short and long yields

The figures depict the fitted par-coupon curves and the observable yields for three days in the sample: January 3, 2005, September 15 2008 and February 23 2015. The left panels show the short end of the yield curve with bonds of less than 20 years to maturity, whilst the right panel shows maturities longer than 20 years. For these bonds noise is based on the extrapolated curve.

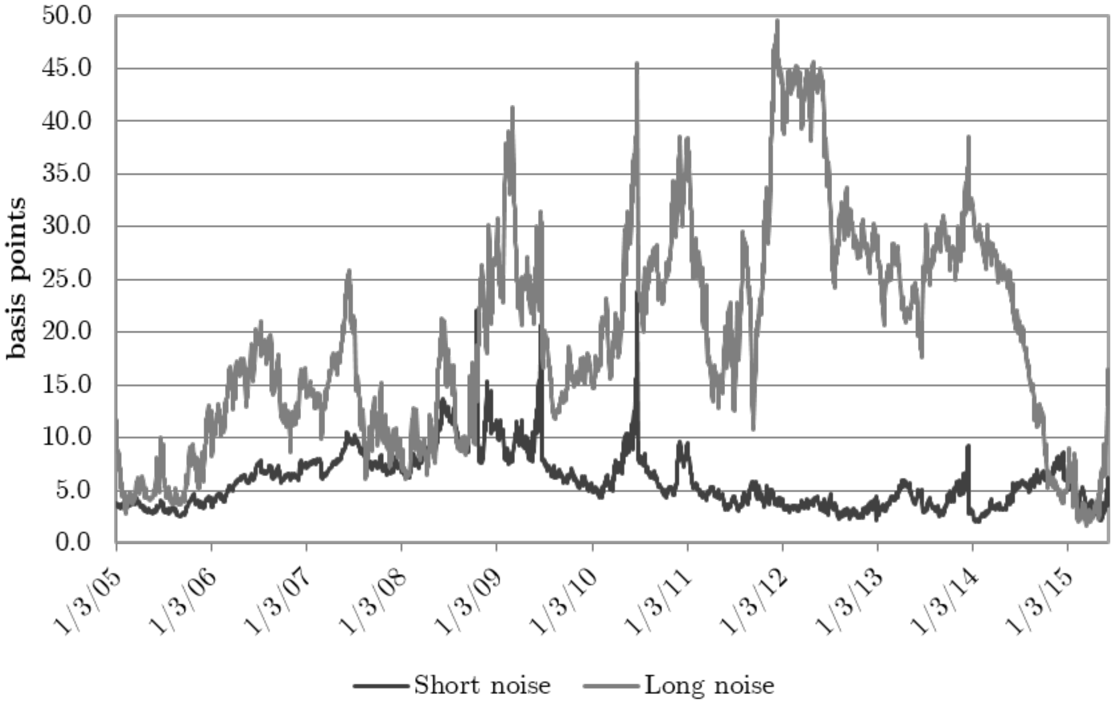

Figure 3.2 Daily timeseries of noise measures in basis points

The figure plots the daily time series of the noise measures for the long and short ends of the yield curve, where long bonds are those of maturities longer than 20 years. Noise is the root mean squared deviation from the smooth Nelson-Siegel curve fitted on German sovereign notes and bonds with maturities between 6 months and 20 years. For bonds with longer maturities, noise is based on the extrapolation of this smooth curve.

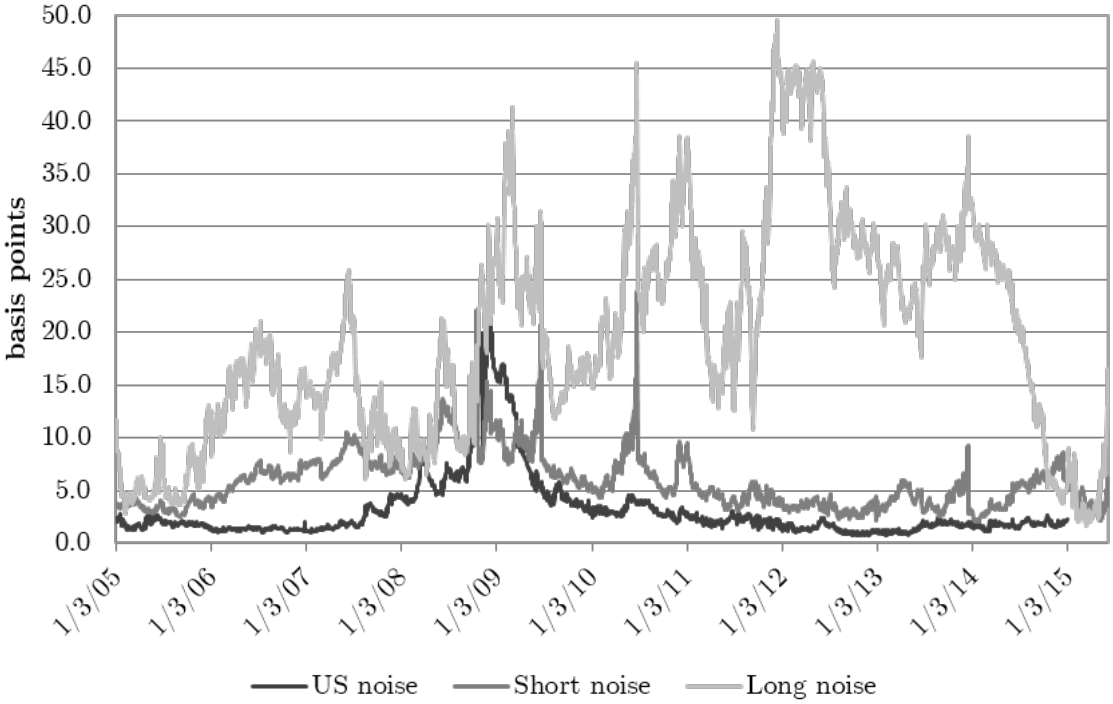

Figure 3.3 Comparison of German and US noise measures

The figure compares the time-series of noise measures of the short and long ends of the German yield curves to the noise measure fitted on US Treasuries by Hu et al. (2013).

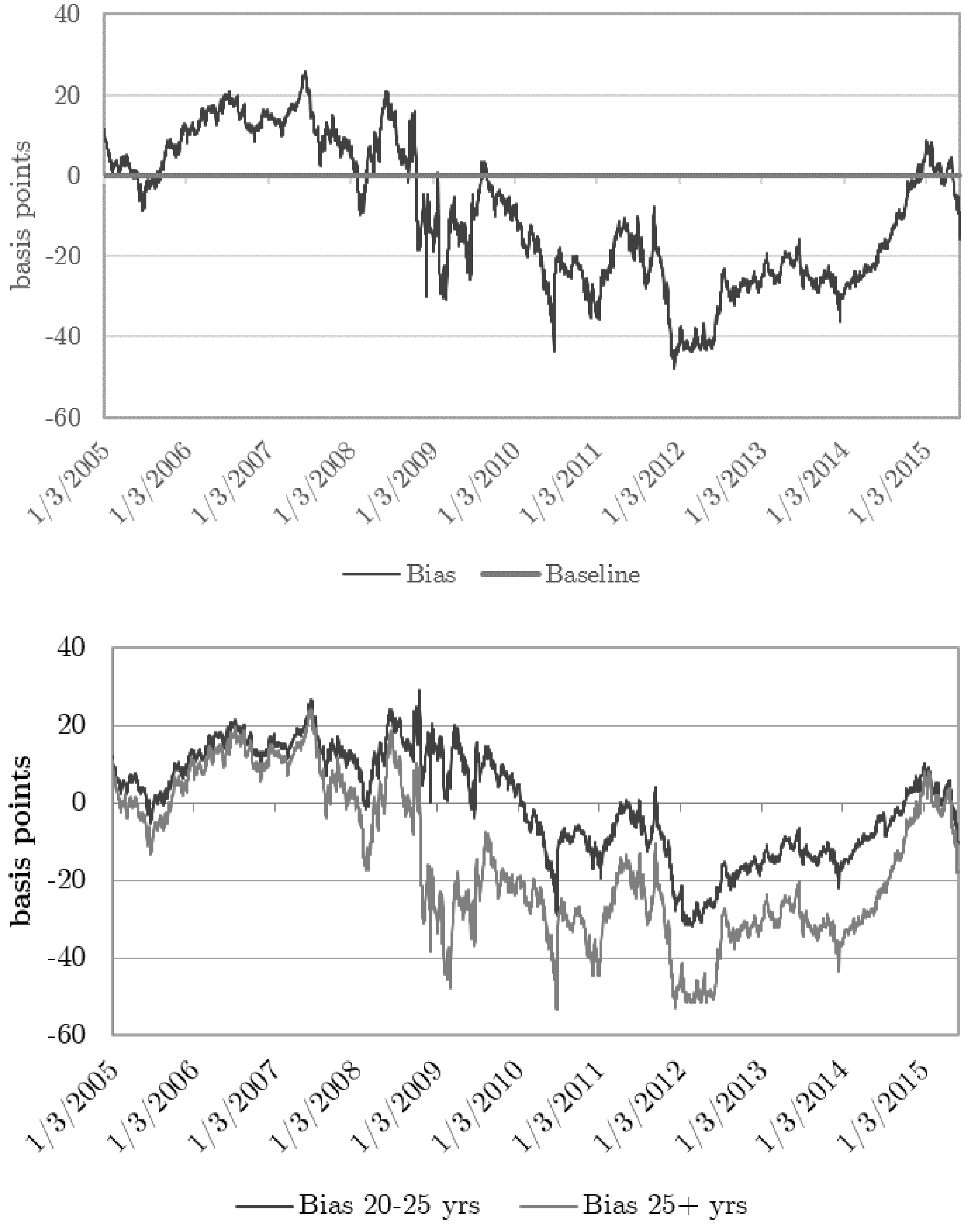

Figure 3.4 Daily time series of the bias

The upper panel of the figure plots the daily time series of the bias. The bias is the structural deviation of long maturity bonds, measured as the fitting error of long maturity bonds relative to the smooth Nelson-Siegel curve, which is fitted on German sovereign notes and bonds with maturities between 6 months and 20 years. The lower panel plots these fitting errors for two maturity buckets: 20-25 and 25+ years to maturity.

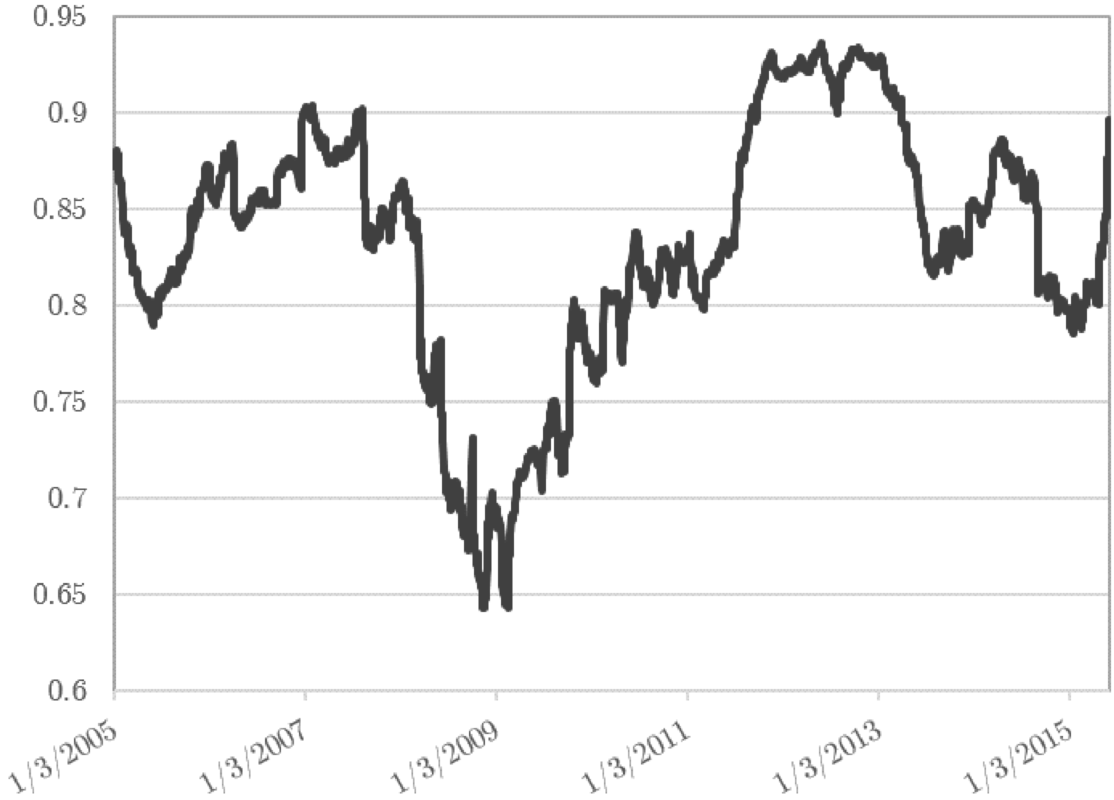

Figure 3.5 Correlation between the change in short and long yields

The figure plots the correlation between first differences of short and long maturity bonds yields, estimated with a 180-day rolling window.

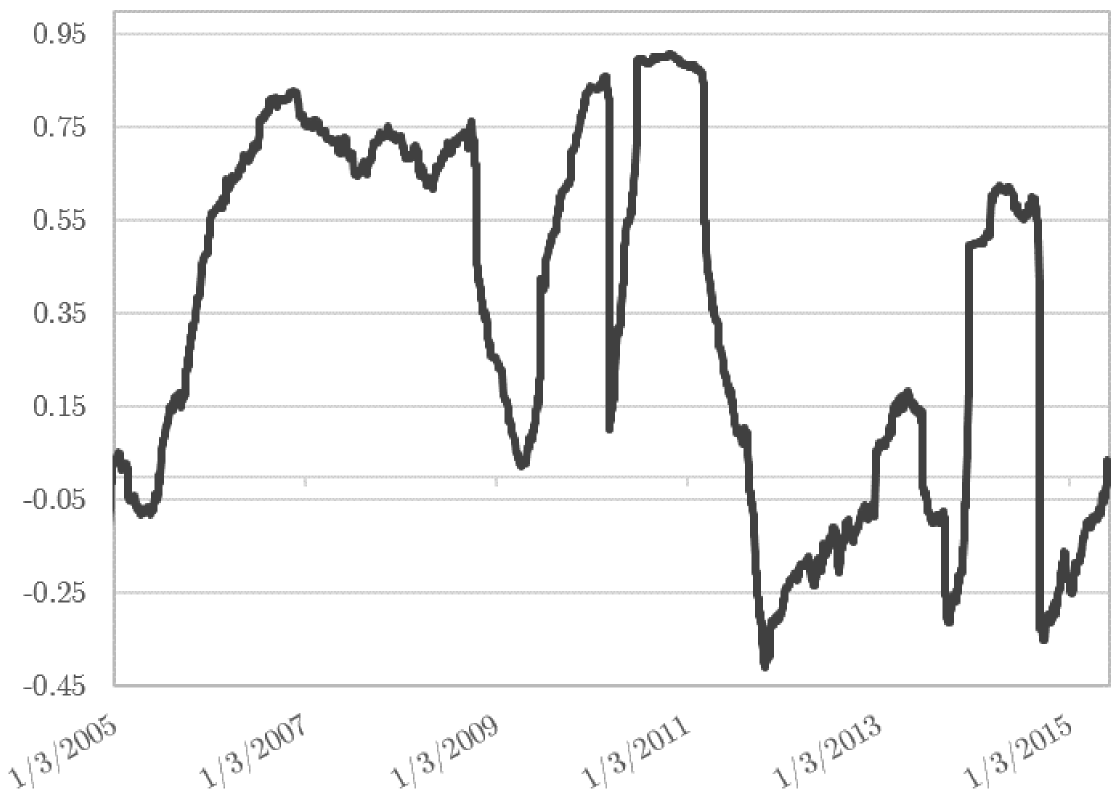

Figure 3.6 Correlation between the change of short and long noise

The figure plots the correlation between the first differences in noise measures of short and long maturity bonds, estimated with a 180-day rolling window.

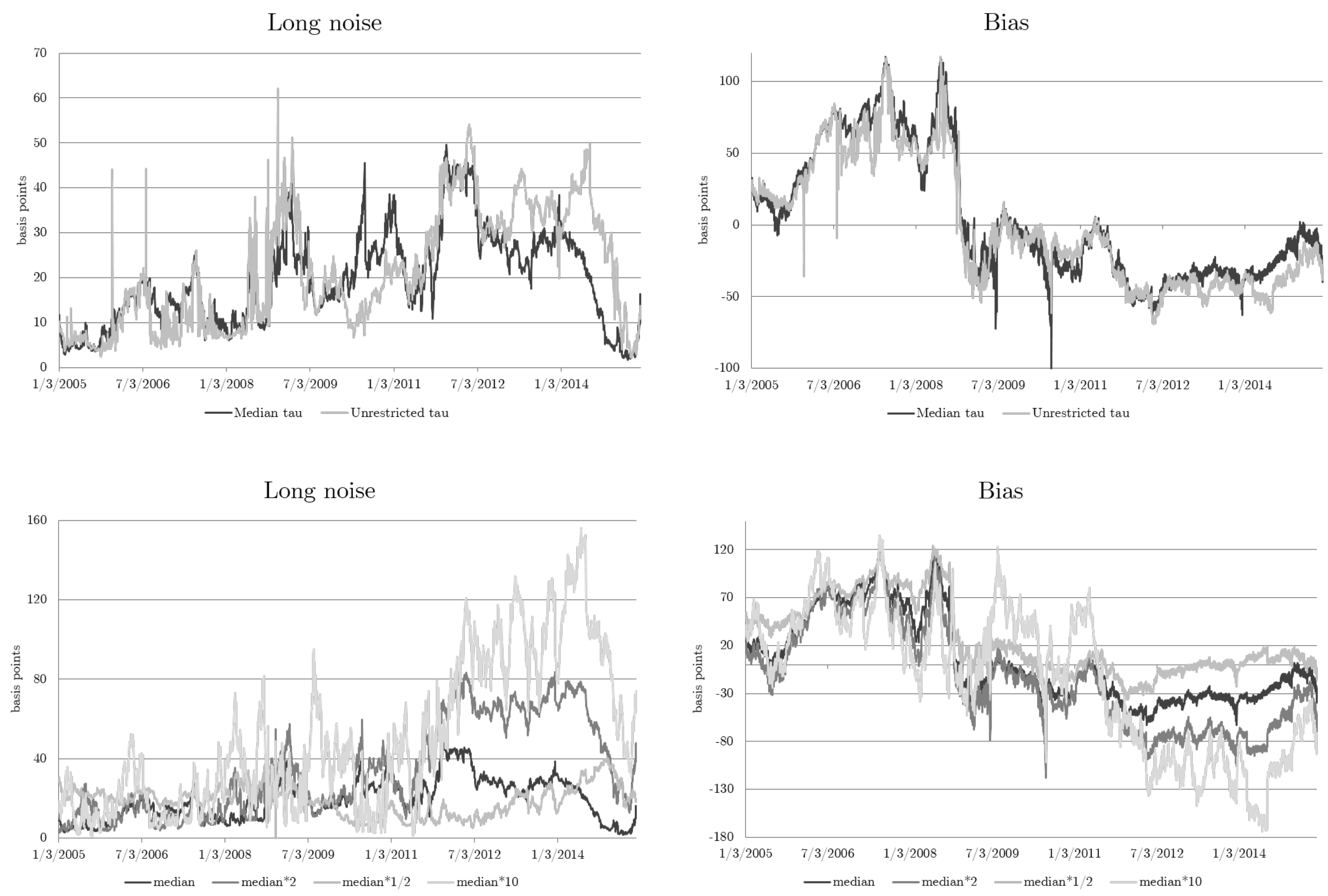

Figure 3.7 Robustness to the choice of τ

The four panels of the figure depict how the choice of tau affects the noise and bias measures of long maturity bonds. We consider different τ parameters from unrestriced and constrained optimiaztion of the Nelson-Siegel curve. The two upper graphs compare the long maturity noise measure and bias based on unrestricetd estimation and one, where taus is fixed at its median value based on the previous method. The lower panel compares noise and bias measures for different fixed values of τ: half the median, median, twice and ten times the unconditional median of the parameter.

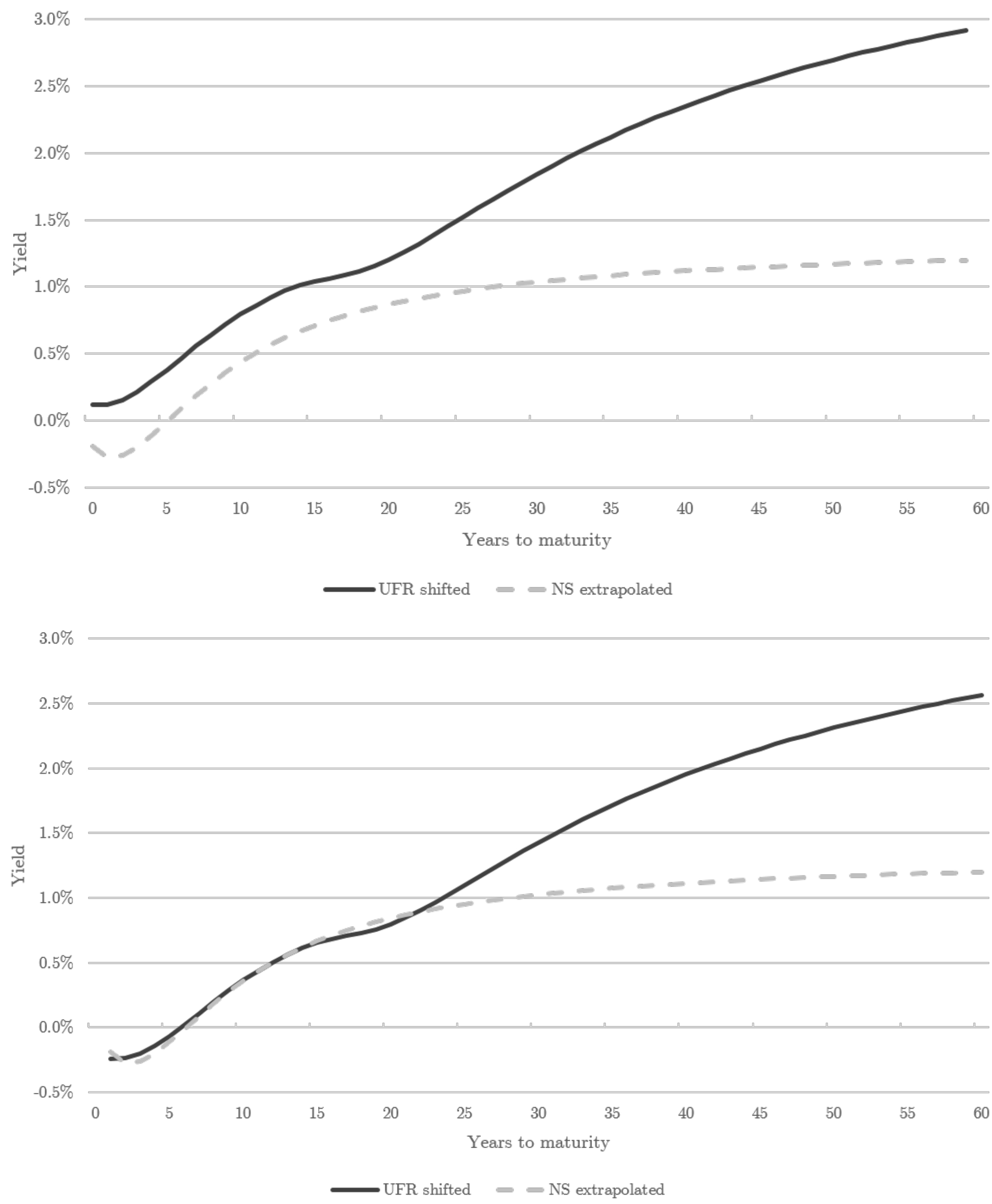

Figure 3.8 Regulatory vs. Nelson Siegel curves

The figures above depict the yield curve on February 23, 2015. Both panels compare the UFR curve, fitted on in interest swap data, to the Nelson Siegel curve of the study extrapolated beyond 20 years to maturity. The panel above depicts the UFR curve as it has been provided by the Dutch Central Bank (DNB, while the panel below shows the unchanged Nelson Siegel curve together with the UFR curve net of the swap premium. The swap premium likely reflects the sum of compensation for counterparty risk in bilateral swap transactions and illiquidity of certain contracts.

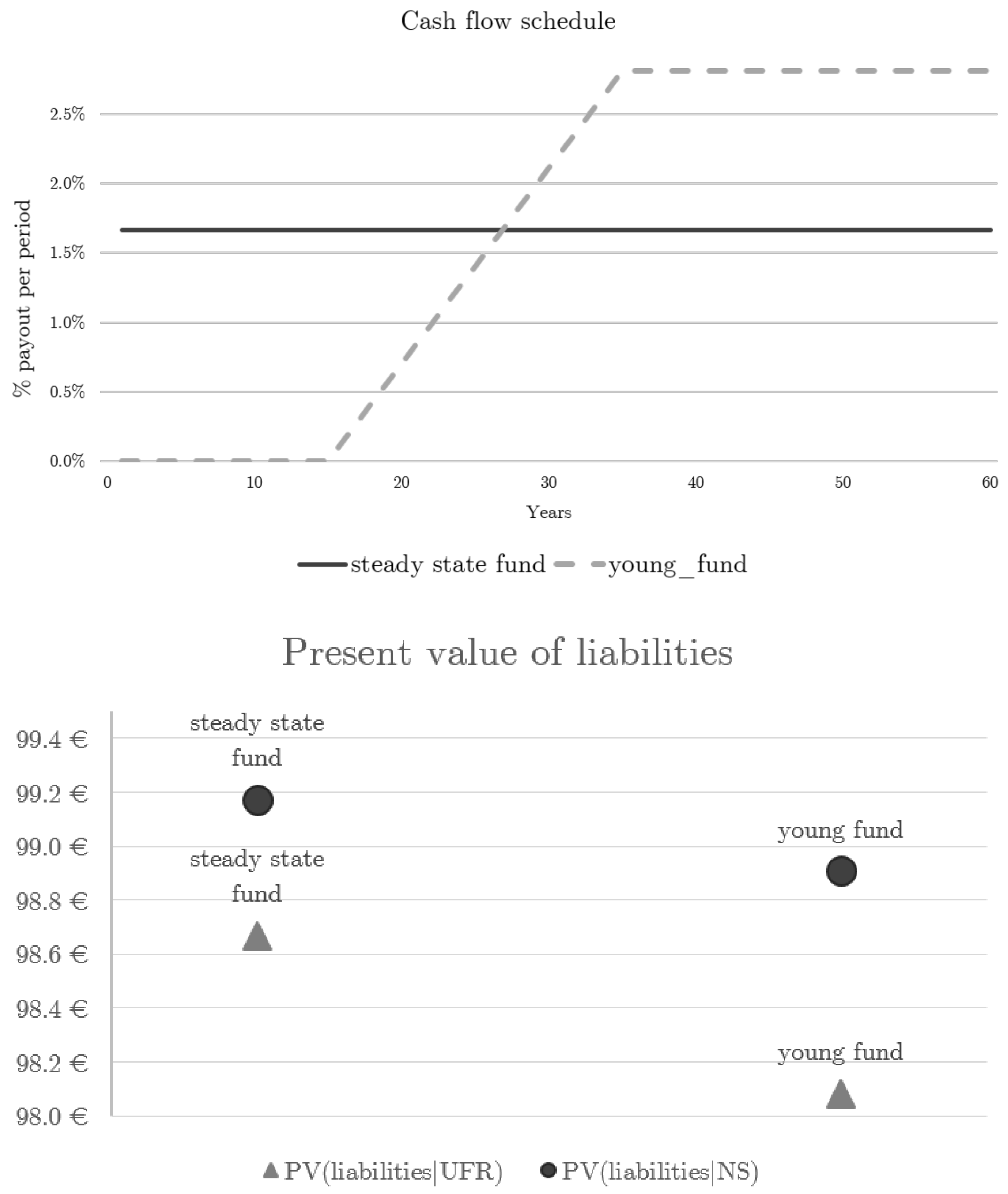

Figure 3.9 Welfare effects of liability valuation: a thought experiment

The figure depicts the liability payout schedules and present values of these liabilities for two hypothetical pension funds: a large and a small and young fund. The upper panel depicts the cash flow structures assumed for the two pension funds in the calculation of welfare effect when one switches from the current regulatory curve to the one we propose for liability valuation. The lower panel compares the present values of liabilities of the two hypothetical funds, where liabilities are either discounted by the shifted UFR or by the extrapolated Nelson Siegel curves.

Bibliography

Acharya, V. and Pedersen, L. H. (2005). Asset Pricing with Liquidity Risk. *Journal of Financial Economics*, 77 (2): 375–410.

Amihud, Y. (2002). Illiquidity and Stock Returns: Cross-Section and Time-Series Effects. *Journal of Financial Markets*, 5 (1): 31–56.

Amihud, Y. and Mendelson, H. (1986). Asset Pricing and the Bid-Ask Spread. *Journal of Financial Economics*, 17 (2): 223–249.

Ang, A., Liu, J., and Schwarz, K. (2008). Using Stocks or Portfolios in Tests of Factor Models. Working paper, Columbia University.

Arce, O., Mayordomo, S., and Pane, J. I. (2011). Do Sovereign CDS and Bond Markets Share the Same Information to Price Credit Risk? Working paper, An Empirical Application to the European Monetary Union Case.

Ashcraft, A., Gârleanu, N., and Pedersen, L. H. (2010). Two Monetary Tools: Interest Rates and Haircuts. Working paper, New York University.

Baele, L., Bekaert, G., Inghelbrecht, K., and Wei, M. (2015). Flights-to-Safety. Working paper 230, National Bank of Belgium.

Bai, J., Julliard, C., and Yuan, K. (2012). Eurozone Sovereign Bond Crisis: Liquidity or Fundamental Contagion? Working paper, Federal Reserve Bank of New York.

Barro, R. J. (2006). Rare Disasters and Asset Markets in the Twentieth Century. *Quarterly Journal of Economics*, 121 (3): 823–866.

Beber, A., Brandt, M. W., and Kavajecz, K. A. (2009). Flight-to-Quality or Flight-to-Liquidity? Evidence from the Euro-Area Bond Market. *Review of Financial Studies*, 22 (3): 925–957.

Beber, A., Driessen, J., and Tuijp, P. (2012). Pricing Liquidity Risk with Heterogeneous Investment Horizons. Working paper.

Bekaert, G., Harvey, C. R., and Lundblad, C. (2007). Liquidity and Expected Returns: Lessons from Emerging Markets. *Review of Financial Studies*, 20 (6): 1783–1831.

Bliss, R. R. and Fama, E. F. (1987). The Information in Long-Maturity Forward Rates. *American Economic Review*, 77 (4): 680–692.

Bolton, P. and Jeanne, O. (2009). Structuring and Restructuring Sovereign Debt: The Role of Seniority. *Review of Economic Studies*, 76 (3): 879–902.

Bongaerts, D., De Jong, F., and Driessen, J. (2011). Derivative Pricing with Liquidity Risk: Theory and Evidence from the Credit Default Swap Market. *Journal of Finance*, 66 (1): 202–240.

Brunnermeier, M. and Pedersen, L. H. (2009). Market Liquidity and Funding Liquidity. *Review of Financial Studies*, 22 (6): 2201–2238.

Bühler, W. and Vonhoff, V. (2011). Term Structures of Liquidity Premia in the U.S. Treasury Market. Working paper.

Buraschi, A. and Jitsov, A. (2005). Inflation Risk Premia and the Expectation Hypothesis. *Journal of Financial Economics*, 75 (2): 429–490.

Calice, G., Chen, J., and Williams, J. M. (2011). Liquidity Spillovers in Sovereign Bond and CDS Markets: An Analysis of The Eurozone Sovereign Debt Crisis. Research paper 2011-105, Paolo Baffi Centre.

Campbell, J. Y., Shiller, R. J., and Viceira, L. M. (2009). Understanding Inflation-Linked Bond Markets. In Romer, D. and Wolfers, J., editors, *Brookings Papers on Economic Activity*, 79–120. Brookings Institution Press.

Campello, M., Chen, L., and Zhang, L. (2008). Expected Returns, Yield Spreads, and Asset Pricing Tests. *Review of Financial Studies*, 21 (3): 1297–1338.

Chen, L., Lesmond, D. A., and Wei, J. (2007). Corporate Yield Spreads and Bond Liquidity. *Journal of Finance*, 62 (1): 119–149.

Chordia, T., Roll, R., and Subrahmanyam, A. (2001). Market Liquidity and Trading Activity. *Journal of Finance*, 56 (2): 501–530.

Christensen, J. H. and Gillan, J. M. (2011). Has the Treasury Benefited from Issuing TIPS? *Federal Reserve Bank of San Francisco Economic Letters*.

Christensen, J. H. E. and Gillan, J. M. (2013). Does Quantitative Easing Affect Market Liquidity? Working paper 2013-26, Federal Reserve Bank of San Francisco.

Ciccarelli, M. and Garcia, J. M. (2009). What Drives Euro Area Breakeven Inflation Rates? Working paper 996, ECB.

Cochrane, J. H. (2005). *Asset Pricing*. Princeton University Press, New Jersey, New Jersey, revised edition.

Cochrane, J. H. and Piazzesi, M. (2002). The Fed and Interest Rates: A High-Frequency Identification. *American Economic Review P&P*, 92 (2): 90–95.

Cochrane, J. H. and Piazzesi, M. (2005). Bond Risk Premia. *American Economic Review*, 95 (1): 138–160.

D'Amico, S., English, W., Lopez-Salido, D., and Nelson, E. (2012). The Federal Reserve's Large-Scale Asset Purchase Programs: Rationale and Effects. Finance and Economics Discussion Series, Working paper 2012-85, Federal Reserve Board.

D'Amico, S., Kim, D. H., and Wei, M. (2010). Tips from TIPS: The Informational Content of Treasury In?ation-Protected Security Prices. Finance and Economics Discussion Series 2010-19, Federal Reserve Board.

D'Amico, S. and King, T. B. (2013). The Flow and Stock Effects of Large-Scale Treasury Purchases: Evidence on the Importance of Local Supply. *Journal of Financial Economics*, 108 (2): 275–564.

Darbha, M. and Dufour, A. (2014). The Term Structure of Bond Market Illiquidity and Default Risk. Working paper.

Daves, P. R. and Ehrhardt, M. C. (1993). Liquidity, Reconstitution, and Value of U.S. Treasury Strips. *Journal of Finance*, 48 (1): 315–329.

De Pooter, M. (2007). Examining the Nelson-Siegel Class of Term Structure Models. Discussion paper 07-043/4, Tinbergen Institute.

De Pooter, M., Martin, R. F., and Pruitt, S. (2013). The Liquidity Effects of Official Bond Market Intervention. Working paper.

De Santis, R. A. (2015). A Measure of Redenomination Risk. Working Paper 1785, ECB.

Dick-Nielsen, J., Feldhütter, P., and Lando, D. (2012). Corporate Bond Liquidity Before and After the Onset of the Subprime Crisis. *Journal of Financial Economics*, 103 (3): 471–492.

Diebold, F. X. and Li, C. (2006). Forecasting the Term Structure of Government Bond Yields. *Journal of Econometrics*, 130 (2): 337–364.

Driessen, J., Nijman, T., and Simon, Z. (2014). The Missing Piece of the Puzzle: Liquidity Premiums in Inflation-Indexed Markets. Discussion paper 02/2014-066, Netspar.

Duffie, D. (2010). Asset Price Dynamics with Slow-Moving Capital. *Journal of Finance*, 65 (4): 1238–1268.

Duffie, D., Pedersen, L. H., and Singleton, K. J. (2003). Modeling Sovereign Yield Spreads: A Case Study of Russian Debt. *Journal of Finance*, 58 (1): 119–159.

Ejsing, J., Grothe, M., and Grothe, O. (2012). Liquidity and Credit Risk Premia in Government Bond Yields. Working paper 1440, ECB.

Ericsson, J. and Renault, O. (2006). Liquidity and Credit Risk. *Journal of Finance*, 61 (5): 2219–2250.

Eser, F. and Schwaab, B. (2016). Evaluating the Impact of Unconventional Monetary Policy Measures: Empirical Evidence from the ECB's Securities Markets Programme. *Journal of Financial Economics*, 119 (1): 147–167.

Fama, E. and French, K. (1993). Common Risk Factors in Returns on Stocks and Bonds. *Journal of Financial Economics*, 33 (1): 3–56.

Fama, E. F. and MacBeth, J. D. (1973). Risk, Return, and Equilibrium: Empirical Tests. *Journal of Political Economy*, 81 (3): 607–636.

Fleckenstein, M. (2013). The Inflation-Indexed Bond Puzzle. Working paper.

Fleckenstein, M., Longstaff, F. A., and Lustig, H. (2014). The TIPS-Treasury Bond Puzzle. *Journal of Finance*, 69 (5): 2151–2197.

Fleming, M. and Sporn, J. (2012). An Analysis of OTC Interest Rate Derivatives Transactions: Implications for Public Reporting. Staff report 557, Federal Reserve Bank of New York.

Fleming, M. J. (2003). Measuring Treasury Market Liquidity. *Federal Reserve Bank of New York Economic Policy Review*, 83–108.

Fleming, M. J. and Krishnan, N. (2012). The Microstructure of the TIPS Market. *Federal Reserve Bank of New York Economic Policy Review*, 27–45.

Fontaine, J.-S. and Garcia, R. (2012). Bond Liquidity Premia. *Review of Financial Studies*, 25 (4): 1207–1254.

Fontana, A. and Scheicher, M. (2010). An Analysis of Euro Area Sovereign CDS and Their Relation with Government Bonds. Working paper 1271, ECB.

Gehde-Trapp, M., Schuster, P., and Uhrig-Homburg, M. (2016). A Heterogeneous Agents Equilibrium Model for the Term Structure of Bond Market Liquidity. Working paper.

Gennaioli, N., Martin, R., and Rossi, S. (2014). Sovereign default, domestic banks, and financial institutions. *The Journal of Finance*, 69 (2): 819–866.

Ghysels, E., Sinko, A., and Valkanov, R. (2007). MIDAS Regressions: Further results and New Directions. *Econometric Reviews*, 26 (1): 53–90.

Gorton, G. B. and Metrick, A. (2010). Haircuts. *Federal Reserve Bank of St. Louis Review*, 507–520.

Goyenko, R., Subrahmanyam, A., and Uhkov, A. (2011). The Term Structure of Bond Market Liquidity and Its Implications for Expected Bond Returns. *Journal of Financial and Quantitative Analysis*, 46 (1): 111–139.

Greenwood, R., Hanson, S. G., and Stein, J. C. (2015). A Comparative-Advantage Approach to Government Debt Maturity. *Journal of Finance*, 70 (4): 1683–1722.

Greenwood, R. and Vayanos, D. (2010). Price Pressure in the Government Bond Market. *American Economic Review*, 100 (2): 585–90.

Greenwood, R. and Vayanos, D. (2014). Bond Supply and Excess Bond Returns. *Review of Financial Studies*, 27 (3): 663–713.

Grishchenko, O. V. and Huang, J. (2013). Inflation Risk Premium: Evidence from the TIPS market. *Journal of Fixed Income*, 22 (4): 5–30.

Gromb, D. and Vayanos, D. (2002). Equilibrium and Welfare in Markets with Financially Constrained Arbitrageurs. *Journal of Financial Economics*, 66 (2): 361–407.

Guibaud, S., Nosbusch, Y., and Vayanos, D. (2013). Bond Market Clienteles, the Yield Curve and the Optimal Maturity Structure of Government Debt. *Review of Financial Studies*, 26 (8): 1914–1961.

Gürkaynak, R. S., Sack, B., and Wright, J. H. (2007). The U.S. Treasury Yield Curve: 1961 to the Present. *Journal of Monetary Economics*, 54 (8): 2291–2304.

Gürkaynak, R. S., Sack, B., and Wright, J. H. (2010). The TIPS Yield Curve and Inflation Compensation. *American Economic Journal: Macroeconomics*, 2 (1): 70–92.

Haubrich, J., Pennacchi, G., and Ritchken, P. (2012). Inflation Expectations, Real Rates, and Risk Premiums: Evidence from Inflation Swaps. *Review of Financial Studies*, 25 (5): 1588–1629.

Houweling, P., Mentink, P., and Vorst, T. (2005). Comparing Possible Proxies of Corporate Bond Liquidity. *Journal of Banking & Finance*, 29 (6): 1331–1358.

Hu, X., Pan, J., and Wang, J. (2013). Noise as Information for Illiquidity. *Journal of Finance*, 68 (6): 2223–2772.

Jordan, B., Jorgensen, R., and Kuipers, D. (2000). The Relative Pricing of U.S. Treasury STRIPS. *Journal of Financial Economics*, 56 (1): 89–123.

Kempf, A., Korn, O., and Uhrig-Homburg, M. (2012). The Term Structure of Illiquidity Premiums. *Journal of Banking & Finance*, 36 (5): 1381–1391.

Kerkhof, J. (2005). Inflation Derivatives Explained. *Fixed Income Quantitative Research, Lehman Brothers*, 1–80.

Klinger, S. and Lando, D. (2015). Safe-Haven CDS Premiuma. Working paper.

Korajczyk, R. A. and Sadka, R. (2008). Pricing the Commonality Across Alternative Measures of Liquidity. *Journal of Financial Economics*, 87 (1): 45–72.

Krishnamurthy, A. (2002). The Bond/Old Bond Spread. *Journal of Financial Economics*, 66 (2): 463–506.

Krishnamurthy, A., Nagel, S., and Vissing-Jorgensen, A. (2015). ECB Policies Involving Government Bond Purchases: Impact and Channels. Working paper.

Krishnamurthy, A. and Vissing-Jorgensen, A. (2010). The Aggregate Demand for Treasury Debt. Working paper, Northwestern University.

Krishnamurthy, A. and Vissing-Jorgensen, A. (2011). The Effects of Quantitative Easing on Interest Rates: Channels and Implications for Policy. Working paper 17555, NBER.

Liu, J., Longstaff, F. A., and Mandell, R. E. (2006). The Market Price of Risk in Interest Rate Swaps: The Roles of Default and Liquidity Risks. *The Journal of Business*, 79 (5): 2337–2360.

Longstaff, F. A. (2004). The Flight to Liquidity Premium in U.S. Treasury Bond Prices. *Journal of Business*, 77 (3): 511–526.

Mitchell, M., Pedersen, L. H., and Pulvino, T. (2007). Slow Moving Capital. *American Economic Review P&P*, 97 (2): 215–220.

Monfort, A. and Renne, J.-P. (2014). Decomposing Euro-Area Sovereign Spreads: Credit and Liquidity Risks. *Review of Finance*, 18 (6): 2103–2151.

Næs, R., Skjeltorp, J. A., and Ødegaard, B. A. (2011). Stock Market Liquidity and the Business Cycle. *The Journal of Finance*, 66 (1): 139–176.

Palladini, G. and Portes, R. (2011). Sovereign CDS and Bond Pricing Dynamics in the Euro-Area. Working paper 17586, NBER.

Pastor, L. and Stambaugh, R. F. (2003). Liquidity Risk and Expected Stock Returns. *Journal of Political Economy*, 111 (3): 642–685.

Pelizzon, L., Subrahmanyam, M., Tomio, D., and Uno, J. (2016 forthcominga). Sovereign Credit Risk, Liquidity, and ECB Intervention. *Journal of Financial Economics*.

Pelizzon, L., Subrahmanyam, M., Uno, J., and Tomio, D. (2014). The Microstructure of the European Sovereign Bond Market. Working paper, New York University.

Pelizzon, L., Subrahmanyam, M. G., Tomio, D., and Uno, J. (2016 forthcomingb). Sovereign Credit Risk, Liquidity, and ECB Intervention. *Journal of Financial Economics*.

Petersen, M. A. (2009). Estimating Standard Errors in Finance Panel Data Sets: Comparing Approaches. *Review of Financial Studies*, 22 (1): 435–480.

Pflueger, C. E. and Viceira, L. M. (2011). Inflation-Indexed Bonds and the Expectations Hypothesis. *Annual Review of Financial Economics*, 3 (1): 139–158.

Pflueger, C. E. and Viceira, L. M. (2015). Return Predictability in the Treasury Market: Real Rates, Inflation, and Liquidity. In Veronesi, P., editor, *Handbook of Fixed-Income Securities*. Wiley and Sons, New Jersey, New Jersey.

Phillips, A. L. (2003). In the Long End of the Spectrum. *Journal of Academy of Business and Economics*, 1 (1).

Quaedvlieg, R. and Schotman, P. C. (2016). Score-Driven Nelson Siegel: Hedging Long-Term Liabilities. Technical report.

Reinhart, C. M. and Rogoff, K. S. (2009). *This Time Is Different: Eight Centuries of Financial Folly*. Number 8973 in Economics Books. Princeton University Press.

Roll, R. (1984). A Simple Implicit Measure of the Bid-Ask Spread in an Efficient Market. *Journal of Finance*, 39 (4): 1127–1139.

Ruenzi, S., Ungeheuer, M., and Weigert, F. (2016). Extreme Downside Liquidity Risk. Working paper 2013/26, University of St. Gallen.

Schuster, P. and Uhrig-Homburg, M. (2013). The Term Structure of Bond Market Liquidity Conditional on the Economic Environment: An Analysis of Government Guaranteed Bonds. Working paper.

Schwarz, K. (2015). Mind the Gap: Disentangling Credit and Liquidity Risks in Spreads. Working paper.

Shleifer, A. and Vishny, R. W. (1997). The Limits of Arbitrage. *Journal of Finance*, 52 (1): 35–55.

Simon, Z. (2015). Not Risk Free: The Relative Pricing of Euro Area Inflation-Indexed and Nominal Bonds. Discussion paper 11/2015-074, Netspar.

Tang, D. and Yan, H. (2007). Liquidity and Credit Default Swap Spreads. Working paper, University of Hong Kong.

Vayanos, D. (2004). Flight to Quality, Flight to Liquidity, and the Pricing of Risk. Working paper, London School of Economics.

Vayanos, D. and Vila, J.-L. (2009). A Preferred-Habitat Model of the Term Structure of Interest Rates. Discussion paper 7547, CEPR.